Provider Supply and Access to Primary Care

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Abstract

Resource-constrained delivery systems often have access issues, in which patients wait a long time to see a provider. We develop theoretical and empirical models of wait times and apply them to primary care delivery by the U.S. Veterans Health Administration (VHA). Using instrumental variables to handle simultaneity issues with supply and demand, we estimate the effect of a clinician supply shock on new patient wait times. We find that one additional clinician per 5,000 enrollees reduced wait times by 3.8 days on average. The reduction varied geographically, ranging from 2.0 to 7.7 days, depending on the local elasticity of demand. The elasticity was affected by local affluence and access to additional health coverage. The VHA has adopted our models to achieve certain goals of the MISSION Act of 2018. Resource-constrained public delivery systems can use this information to prioritize hiring clinicians in areas yielding the highest access return.

Keywords: wait time; access to care; resource allocation; public health care; provider supply capacity; demand elasticity

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1. INTRODUCTION

In many countries, patients must wait a long time to receive health care (Merritt Hawkins 2017; Viberg et al. 2013). Nicolova et al. (2016), Sutherland et al. (2016), and Pizer and Prentice (2011) document the potential adverse effects of waiting on patient outcomes. Access issues are exacerbated in supply-constrained systems, such as public health care delivery systems, and managing such issues is not trivial. Reducing demand through cost sharing is unpopular (Siciliani and Hurst, 2005). Reducing wait times by increasing capacity or improving productivity requires additional resources. For many systems, a better approach may be to reevaluate the allocation of resources and target investments to achieve maximum impact.

We develop a supply and demand model that illustrates how resources could be reallocated to reduce wait times more efficiently. Empirically, we apply our model to the U.S. Veterans Health Administration (VHA), the largest public delivery system in the U.S. We examine factors that affect the wait times for new patients seeking primary care appointments at the VHA. Using a fixed-effects specification and instrumental variables, we investigate geographic variation in the responsiveness of wait times with respect to shocks in clinician supply. We demonstrate (through a simulation exercise) how our results can be used to target investments to improve access.

Our conceptual model builds off a substantial literature starting with Lindsay and Feigenbaum (1984) who were the first to show that wait times are affected by changes in patient demand for health care services. Authors in this tradition assume that the value of care decays as time passes between diagnosis and treatment (e.g., Goddard et al., 1995; Besley et al., 1999; Martin and Smith, 1999; Goddard and Smith, 2001). Consequently, the waiting time acts like a price – imposing a cost on the patient that reduces demand. The focus has largely been on the demand for health care, so previous studies typically assume that supply is positively related to wait times because managers are altruistic or incentivized to care about wait times. We revisit this assumption and show that supply, in fact, is also mechanically related to wait times. While this does not translate to different qualitative implications, it clarifies the supply side of the model and supports applications.

Empirically, we contribute to the literature in several ways. First, we directly estimate the effect of an increase in clinician supply on wait times. A set of empirical studies have estimated demand and supply elasticities, thereby showing the relationship between wait times and utilization. Cullis et al.

(2000) and Siciliani and Iversen (2012) provide excellent reviews of the literature. While these estimates are useful in forecasting access to care issues[5] (and can be used to perform back-of-the-envelope calculations of how a shift in supply affects wait times), our direct estimate of the *capacity effect* is more operationally useful for managers of health care systems. Another set of studies have examined natural experiments that shift supply. Siciliani and Martin (2007) and Propper et al. (2008) examined changes in hospital competition; Propper et al. (2002) and Dusheiko et al. (2004) examined a national (U.K.) policy change in provider incentive structures; Propper et al. (2010) examined another national (U.K.) policy that forced hospitals to have wait times below a certain threshold. We estimate the impact of a different type of supply measure – the clinician workforce – on wait times.

Second, we broaden the scope of the previous literature, which has largely focused on elective surgery in national delivery systems in European countries, i.e. the U.K., Italy, Australia, and Norway (e.g., Martin and Smith, 1999, 2003; Gravelle et al., 2002, 2003; Martin et al., 2007; Fabbri and Monfardini, 2009; Riganti et al., 2017; Stavrunova & Yerokhin, 2011; Monstad et al., 2006; Sivey, 2012). To our knowledge, supply and demand for primary care services has not been studied previously. Moreover, the responsiveness of wait times with respect to supply shocks has not been estimated for the U.S., where the delivery and financing of health care is different in many ways from the pattern seen in other countries (Mossailos et al., 2016).

Finally, we document the geographic variation in the effect of clinician capacity on wait times. We analyze which local factors (e.g., affluence, economic indicators, and access to alternative health insurance) influence the magnitude of the capacity effect, i.e., the sensitivity of wait times with respect to an increase in clinician capacity. We find that the variation is large and dependent on local income levels and (depending on specification) access to alternative health insurance. The capacity effect is larger in areas with lower income and less access to other health insurance. This suggests that delivery systems should consider strategically allocating resources – prioritizing certain areas over others – in order to maximize access (and population health) under financial constraint.

This study is timely for U.S. policy. In 2014, several veterans died while waiting to see a VHA provider, highlighting the access issue for veterans. In response, the Veterans Choice Act of 2014 and

[5] Supply elasticities help to predict how much wait times will lengthen due to an outward shift in patient demand. Demand elasticities help to predict how much wait times will fall given an outward shift in supply

the MISSION Act of 2018[6] dedicate public VHA funds to pay for privately provided care. While these Acts may increase the number of available clinicians, it is unclear whether paying for private care will improve timeliness of care at an affordable price (Yee et al. 2016). With limited funds, the VHA should make efficient use of spending, identifying the areas that will yield the highest return (greatest access improvement) for the next dollar spent.

Indeed, one of the goals of the MISSION Act (Title IV Section 401) is to identify underserved areas and create a strategic plan to address the access issue in these areas. Since our model has this exact application, it has been adopted by the VHA to comply with and reach this goal of the MISSION Act. We discuss this policy application in the Discussion Section.

2. METHODOLOGY

Conceptual Model

In a cash market in which patients pay the full price of health care, excess demand drives up prices until equilibrium is reached. However, in many health care systems prices are established by regulation or contracts and held fixed for a duration longer than demand fluctuations. In such settings, prices cannot rise in the short term. Instead, variation in waiting time through queues take the place of changes in out-of-pocket prices (Goddard et al., 1995). The equilibrium health care quantity delivered is determined by both out-of-pocket price and wait time.

At a given price, the supply and demand curves in the *wait-appointment* space are analogous in many ways to the curves in a *price-appointment* space that hold constant the wait times. As wait times increase (and price is fixed), fewer patients will demand appointments from a given provider (Equation 1).

$$Appt^D = D(wait; alternative\ coverage, affluence, need) \qquad (1)$$

A patient's decision to schedule an appointment is based on his or her willingness to wait, compared to the number of days until the next available appointment, denoted in the equation by *wait*. Willingness to wait for an appointment with a particular provider is dependent on the patient's *alternative health*

[6] The Choice Act is short for the Veterans' Access to Care through Choice, Accountability, and Transparency Act. The MISSION Act is short for Maintaining Systems and Strengthening Integrated Outside Networks Act.

coverage options, which provide access to other providers. Willingness to wait is also dependent on the patient's *affluence*. Having more income gives one more flexibility in choosing a provider. It also may be related to one's value of time, and thus willingness to wait. Finally, the *need* for medical services may influence which provider a patient chooses.

In the case of the VHA, if veterans have additional health care coverage – e.g., Medicaid or Medicare – or have the financial means to pay for such alternative care, they may demand less care from the VHA. Their willingness to wait may be lower, and they may be more sensitive to changes in wait times. If VHA wait times exogenously increased, veterans with other coverage could easily shift to alternative providers and may prefer to do so, despite having little or no copayments for VHA care and potentially more coordinated and higher quality of care at the VHA (Trivedi et al. 2011).

On the supply side, the number of appointments supplied is directly determined by the number of clinicians, how productive each one is, and the wait time. For example, suppose a newly hired clinician can perform five appointments in a day. The first five patients who desired an appointment (patients 1-5) would have a zero-day wait time, patients 6-10 would have a 1-day wait time, and this continues until the wait time is too high for the next patient to desire scheduling an appointment. If the wait time is 9 days, the clinician must be scheduled to provide 50 appointments (10 days x 1 clinician x 5 appointments per day). If a provider group has two clinicians and a 9-day wait time, the number of appointments would be 10 days x 2 clinicians x the average productivity of the two clinicians. This logic holds if we assume that patients always take the first available appointment. Thus, the number of appointments supplied is

$$Appt^S = Wait \times \#\, clinicians(E(wait), z) \times \#appts\ per\ clinician\ per\ day(x) \qquad (2)$$

where *Wait* is the wait time in days at a given point in time *t*; *# appts per clinician per day* is the potential productivity; and *# clinicians* is the full-time equivalent number of clinicians. The number of clinicians on staff is a decision made by health care managers. It likely depends on expected wait times ($E(wait)$) and supply shifters (*z*), such as lower costs to retaining clinicians, changes in infrastructure, and expansion of physical capacity. Productivity can also depend on supply shifters (*x*), such as technology that enables shorter appointment times.

This specification of supply is different from the previous literature (e.g. Martin and Smith, 1999, 2003), which assumed that managers generate more supply when contemporaneous wait times are higher, due to being altruistic or incentivized (financially or otherwise) to improve access. In Equation 2, we distinguish wait from expected wait. Expected wait is in line with the previous literature. Moreover, in the setting of the VHA, supply of clinicians (via budget appropriations) is dependent on forecasted demand and expected wait times. Equation 2 not only exhibits this indirect relationship with wait times through expected wait times, but also the mechanical relationship with wait times. This distinction allows us to consider the direct and indirect avenues through which wait times are related to appointments supplied[7].

Equilibrium occurs when the number of appointments demanded (Equation 1) is the same as that supplied (Equation 2). The equality can be rearranged, such that $D(wait; alternative\ coverage, affluence, need) \times S^{-1}(wait; clinicians, productivity) - 1 = 0$, to which we can apply the implicit function theorem. This allows us to think of wait time as a function of alternative coverage, affluence, need, clinicians, and productivity.

If we are willing to make assumptions regarding functional form of the structural demand and supply functions, we can qualitatively describe the relationship between the effect of a change in the number of clinicians and the underlying structural parameter of the demand elasticity with respect to wait times. For example, let's assume a log-linear additive demand equation, i.e., $\ln(appt^D) = \beta_0^D + \beta_1^D \cdot wait + \beta_2^D \cdot alternative + \beta_3^D \cdot affluence + \beta_4^D \cdot need + \epsilon$. The supply function Equation 2 can be transformed to $\ln(appt^S) = wait + clinicians(E(wait), z) + productivity(x)$[8]. We can further partition clinicians, such that $\ln(appt^S) = wait + clinicians(E(wait)|z) + clinicians(z|wait) + productivity(x)$. Depending on assumptions about the relationship between clinicians and current wait times, we will have different reduced form models.

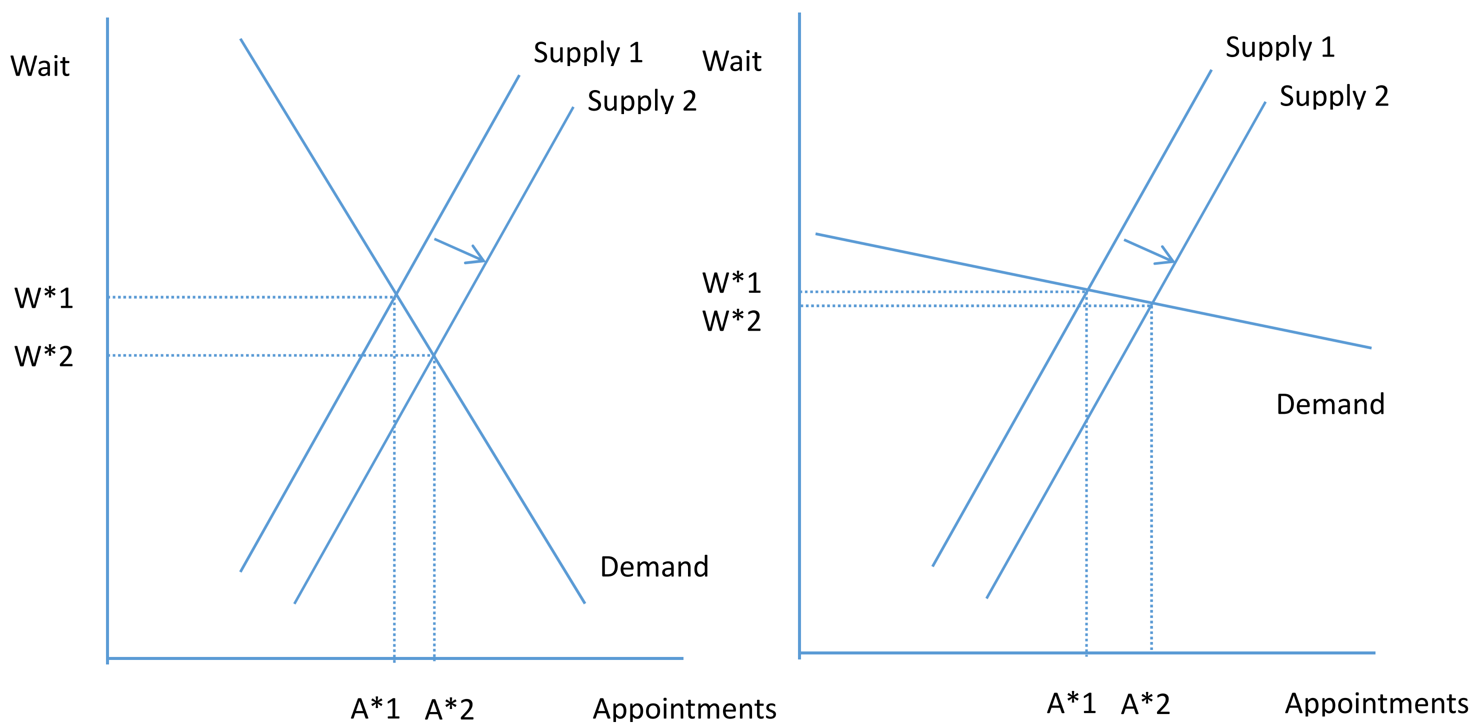
In the simplest case, the number of clinicians is not determined by expected wait times ($clinicians(E(wait)|z) = 0$). Setting $\ln(appt^D) = \ln(appt^S)$, the reduced form equation under either of these assumptions would be

[7] This setup also departs from the traditional supply-demand framework, which uses price as the rationing mechanism. In the traditional framework, price affects supply by incentivizing (for-profit or non-profit) firms to produce more via some sort of optimization process. It is not mechanically related to supply.
[8] Note that there are no coefficients because the equation is true by definition, assuming Equation 2 holds.

$$wait = \frac{\beta_0^D + \beta_2^D \cdot alternative + \beta_3^D \cdot affluence + \beta_4^D \cdot need - clinicians - productivity + \epsilon}{1 - \beta_1^D} \quad (3)$$

This reduced form model shows that our policy lever of interest – the effect of hiring more clinicians on wait times – is inversely related to the elasticity of demand with respect to the wait time. The value $-\frac{1}{1-\beta_1^D}$ is negative if $1 - \beta_1^D$ is positive. This certainly happens in the case $\beta_1^D < 0$, i.e., the demand curve in the wait-appointment space is downward sloping. If demand is less elastic (i.e., the elasticity $\beta_1^D \cdot wait$ is small in magnitude), an additional clinician would reduce wait times by a large amount, assuming all else is held constant (Figure 2-1a). If demand is elastic, the newly available appointments that result from a positive supply shock would be filled quickly by patients, and wait times would not fall by as much as they would have fallen given the same supply shock but in a market with inelastic demand (Figure 2-1b).

Figure 2-1. (a) Inelastic Demand and (b) Elastic Demand for Appointments



Realistically, the number of clinicians may be related to expected wait times ($clinicians(E(wait)|z) \neq 0$), and expected wait times are likely to be a function of current wait times. One way to illustrate this is to assume $clinicians(E(wait)|z) = \alpha \cdot wait$. In this case, the supply equation would be $\ln(appt^S) = (1+\alpha) \cdot wait + clinicians(z|wait) + productivity(x)$, where α represents the relationship between the number of clinicians and the wait time (via the expected wait time). The reduced form equation would be

$$\boldsymbol{wait = \frac{\beta_0^D + \beta_2^D \cdot alternative + \beta_3^D \cdot affluence + \beta_4^D \cdot need - clinicians - productivity + \epsilon}{1 + \alpha - \beta_1^D}} \qquad (3')$$

The effect of hiring a clinician on the wait time $\left(-\frac{1}{1+\alpha-\beta_1^D}\right)$ is negative if $1 + \alpha - \beta_1^D$ is positive, for example, when $\alpha > 0$ and $\beta_1^D < 0$. It will be larger or smaller depending on the sensitivity of managers with respect to the wait time (via expected wait time) and the elasticity of demand. α functions as a parameter that modulates the elasticity of supply. A large and positive α suggests that the hiring decisions of managers are very sensitive to wait times, suggesting that supply is more elastic. A large and negative β_1^D implies that the demand is elastic with respect to wait times. The composite effect of clinician capacity on the wait time will be larger when the supply of clinicians is less responsive to the wait time and the demand for appointments is less responsive to the wait time.

Empirically, we can reduce the influence of the supply component – thereby getting closer to Equation 3 than 3' – by using instruments (*z)* that exogenously affect the number of clinicians. For example, in the case of the VHA, there may be a change in the private health care market that affects the VHA's ability to hire. Equation 3 differs from previous studies' reduced form models (Siciliani and Iversen 2012 Equation 24.3) because clinicians and the wait time enter our structural supply equation with a coefficient equal to 1 (due to the mechanical relationship described by Equation 2). Previous studies also differ because they assumed their supply variables were exogenous, which we do not. In some cases, this may have been valid because the supply variable was the number of beds, which is less variable in the short term than our supply variable of interest – the number of clinicians. Since clinician supply is likely to be simultaneously determined with wait and number of appointments, we use instrumental variables. Due to this estimation strategy, other empirical issues, and the strong assumptions to get to Equation 3, our empirical specification is not Equation 3 exactly. Consequently, we perform additional analyses to compute the elasticity of demand (our central focus), which are explained in the Discussion Section.

Instrumental Variables

To mitigate bias caused by simultaneous changes in demand for VHA care and VHA capacity, we need to exploit shocks to VHA production costs. We test two instrumental variables for VHA clinician capacity: 1) the turnover rate among registered nurses who have worked at the VHA for less than five years, and 2) the number of non-federal specialists per capita. These variables are plausibly correlated with clinician capacity because they reflect the ability of the VHA to hire or retain clinicians and thus VHA production costs. Higher turnover among new registered nurses potentially suggests a worse work environment, and as we will show in the Results Section, is correlated with primary care clinician turnover. The number of non-federal specialists per capita is indicative of the attractiveness of the local area for medical providers. The higher the number, the easier it is for the VHA to hire clinicians. We discuss the strength of the instruments in the Results Section.

Pertaining to the exclusion restriction, the instruments are plausibly not related to VHA wait times for new patient, primary care appointments except through their effect on VHA primary care clinician capacity. Nurse turnover, especially among new nurses, likely does not affect the scheduling of new patient appointments. These decisions are typically based on the schedule of clinicians not nurses. Moreover, we include all new nurses employed at a given VHA medical center not just the ones working in primary care clinics, and we use less-tenured nurses who are likely less integral to the system than senior ones, thus mitigating a direct relationship with primary care wait times. It is plausible, however, that the turnover among new registered nurses is (negatively) correlated with quality, and quality could affect the demand for VHA care, thus impacting (lowering) wait times through a mechanism other than primary care clinician capacity. This is a countering effect to that which works through clinician capacity (high turnover, lower clinician capacity, higher wait times), so it would bias the estimate toward zero.

The number of non-federal specialists per capita is plausibly not directly related to VHA wait times for new patient primary care appointments. Since it characterizes the private health care market – outside the VHA – it is unlikely to be influenced by VHA actions. VHA-provided health care is a small fraction of the health care provided by private providers. Moreover, this variable excludes primary care physicians (in the private market), which further reduces any relationship with wait times for primary care appointments in the VHA.

It is possible, however, that more private specialists equate to lower wait times for non-VHA specialized services, thus lowering the demand for VHA specialty care. If veterans seek VHA primary care to get access to VHA specialty care, the demand for VHA primary care may be less in areas with more private specialists, which would lower the wait times. This effect would be in the same direction as that which works through clinician capacity (more private specialists, more VHA primary clinicians, lower wait times) and thus might bias our estimate to be larger in magnitude. However, if the number of non-VHA specialists is positively related to wait times for non-VHA specialty care (controlling for other factors in the model) and these wait times are positively related to wait times for VHA primary care, our estimate would be biased toward zero. Although we were not able to estimate these correlations, we did estimate the referral rate of new patient veterans to see a VHA specialist. While this is not the same as the proportion who sought VHA primary care with the *intention* of acquiring a referral, it provides an upper bound. We found that 24% of new patient veterans received such a referral, which suggests seeking a referral does not explain most of the reason new patients seek primary care.

Data Sources and Sample

We used multiple administrative data sets for this study. The VA Corporate Data Warehouse is a national repository of several VA clinical and administrative systems. The data contain the date on which a veteran made an appointment and the date of the corresponding clinic visit (which we use to derive wait times). The VHA Enrollee Table contains enrollment status of veterans, the priority status of veterans (see Appendix for more information), veteran age, county of residence, and the VHA facility that is closest to them. Data on each medical centers' clinician capacity was provided by the VHA's Office of Productivity, Efficiency, and Staffing (OPES). OPES has monthly data on the number and workload of primary care clinicians (i.e. internal medicine and geriatric medicine physicians). The VHA's Human Resources Department provided data on the number of nurses, level of nurse qualification (e.g. registered or licensed practical nurse), tenure of nurses, and how many quit each month.

Our study sample includes 127 VHA medical centers in the U.S. We excluded several medical centers that either were not in the U.S. (e.g. Manila, Philippines) or had limited data (e.g., new medical centers). The unit of observation in the analysis is a medical center-year-month. The data span July 2010 to June 2016.

We merged the VHA data to non-VHA data, which describe local economic and health care market conditions. The Area Health and Resource File (AHRF) provides county-year level data on the number of non-federal, active physicians; non-federal, active primary care physicians; household median income level; the proportion of males aged 18-64 who report having health insurance; and population count. Zillow provides a monthly housing price index at the county level (called the Home Value Index). The American Community Survey provides an annual, county-level veteran unemployment rate. The Centers for Medicare and Medicaid Services provides a monthly Medicare Advantage penetration rate by county. Finally, The Kaiser Family Foundation provides annual data on whether a state had expanded Medicaid or not.

We aggregated and merged variables from VHA and non-VHA sources to the VHA medical center level. Some VHA variables pertain to the patients or users of VHA health care and other variables pertain to the enrollees in VHA health care. Not every veteran enrolls in VHA health care, and not every enrolled veteran uses or receives VHA health care. Approximately 9 of the 20 million U.S. veterans are enrolled in VHA health care, and 5 of the 9 million enrollees receive VHA services in a typical year. Enrolled veterans were assigned to the medical center closest to their residential addresses. We linked county-level variables from non-VHA sources to each medical center by taking a weighted average of the values associated with the counties in which enrolled veterans resided (referred to as the center's *catchment area*). The weights were the proportion of a medical center's veteran enrollees (in 2010) who resided in a given county. For VHA and non-VHA data that were only available at the annual level, we interpolated the variables to the month level[9]. We discuss the construction of our variables in more detail in the next section. Appendix Table 6-3 provides more information on the data sources and the level of aggregation of the original data, which we convert to the medical center-year-month level.

[9] The month that was assigned to the values provided by AHRF for a given year depends on the month that the data was collected, as reported in the AHRF documentation. For example, if the survey of the number of primary care physicians was conducted in December of each year, we interpolated monthly values from December of year Y to December of year Y+1. The interpolated values were then aggregated to the facility catchment area-year-month level.

Empirical Model and Variable Construction

We estimate the following reduced form model of primary care wait times for new patients on clinician capacity and control variables.

$$Wait_{f,t} = \beta_1 \cdot \frac{FTE}{enrollee_{f,t}} + \boldsymbol{\delta} \cdot \boldsymbol{Interaction}\left(\frac{FTE}{enrollee_{f,t}}\right) + \beta_2 \cdot Income_{f,t} + \beta_3 \cdot ZHPI_{f,t} + \beta_4 \cdot \%VetCopay_{f,t}$$
$$+\beta_5 \cdot \%VetUnemp_{f,t} + \beta_6 \cdot \%\mathrm{Vet65}_{\mathrm{f,t}} + \beta_7 \cdot \%HealthIns_{f,t} + \beta_8 \cdot MedicaidExp + \beta_9 \cdot \%MedicareAdv + \alpha_f + y_t + m_t + u_{f,t} \tag{5}$$

The model includes measures of clinician supply, alternative health coverage options, affluence, facility fixed effects (α_f), year indicators to control for national trends (y_t) in new patient primary care wait times, and month indicators to control for seasonality (m_t)[10].

The dependent variable is the average wait time for new patient primary care appointments at a given VHA facility (*f*) in a given year-month (*t*)[11]. The wait time is the number of days between when a veteran made an appointment and the scheduled clinic visit date[12]. A new patient is defined as a veteran who has not seen a VHA primary care clinician in two or more years.

The main independent variable of interest is the number of full-time-equivalent primary care clinicians per 1,000 enrollees at each medical center, referred to as *clinician capacity*. Primary care clinicians include gerontology medicine and internal medicine physicians[13]. We normalize the number of

[10] This differs from the reduced form model described above in the way that clinicians is defined – the empirical model scales clinicians by the number of enrollees. We also do not include productivity in the empirical model. The theoretical reduced form model is based on log-linear structural supply and demand functions. In the Discussion Section, we compute the elasticity of demand and compare to the previous literature.
[11] Primary care was identified by three VHA stop codes: 323, 350, and 322.
[12] To avoid issues of misreporting (as documented by investigations following the VA Scandal of 2014), we did not use the time between the desired date of appointment and the appointment date as the wait time measure (Office of Inspector General 2014).
[13] Each Physician person class from VHA Person Class Fields in the local VISTA New Person File is mapped into a specialty that is defined by the American Board of Medical Specialties (ABMS). The ABMS defined specialty is then grouped into an aggregated specialty for reporting in the Physician Productivity cube and OPES Physician Workforce Reports. Active physician person class data captured on outpatient and inpatient encounter based workload that is transmitted to the National Patient Care Database (NPCD) from the local VISTA PCE (Patient Care Encounter) package is summarized under the appropriate ABMS Specialty. Note: A physician's total FTEE and encounter workload, regardless of where the workload occurred, will be assigned to their respective ABMS Specialty.

clinicians by the number of enrollees because one new clinician hire would not have the same impact in a large facility with many clinicians and enrollees as in a small facility[14].

As discussed previously, we do not assume exogeneity of clinician capacity and test two instrumental variables for clinician capacity. The monthly turnover rate of registered nurses who have worked at a given VHA facility for less than five years was provided by VHA Human Resources and is not specific to primary care. The number of non-federal, physician specialists per 1,000 capita in the facility's catchment area was constructed using AHRF data – subtracting the number of non-federal, active primary care physicians from total non-federal, active physicians for each county, dividing by the population in thousands, and weighting up the counties to the catchment area. We perform weak instrument tests for each instrument separately and jointly, and ultimately show results from the permutation that yields the most strength.

In addition to primary care clinician capacity, VHA wait times may be affected by alternative health coverage options, affluence, and economic factors. The economic and affluence measures in the model include the median household income level, the Zillow housing price index[15], veteran unemployment rates, and the proportion of veterans who are subject to cost sharing (i.e., who are not disabled and who often earn more than a certain income level – see Appendix Table 6-1 and Table 6-2). Our measures of the influence of alternative health coverage options include: the proportion of VHA-enrolled veterans who are age-eligible for Medicare, whether the VHA medical center's state expanded Medicaid, the Medicare Advantage penetration rate, and the proportion of males 18 to 64 years old with health insurance. Veteran unemployment rates could also be related to alternative health coverage as a (negatively correlated) proxy for the proportion of veterans with employer-sponsored insurance. We test whether these variables influence the level of wait times. We also test whether some of them affect the responsiveness of wait times to changes in clinician capacity.

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[14] We recognize that by dividing by the number of enrollees rather than the actual caseload of the primary care division, we may be using a less accurate depiction normalized capacity. However, we did not use the division's actual caseload because it is likely to be correlated with wait times and thus suffer from endogeneity issues. In contrast, local enrollment is stable – once enrolled, veterans remain enrolled for life, regardless of whether they use VHA services.

[15] We also tested the Federal Housing Finance Agency's Purchase-Only and All-Transactions Housing Price Indices. However, these measures were geographically sparser than the Zillow Home Value Index, and they only provided quarterly data, whereas the Zillow Home Value Index provided monthly data. The results are qualitatively similar. We can provide results using the Federal Housing Finance Agency Indices upon request.

The coefficient β_1 can be interpreted as the effect of an increase in primary care clinician capacity (via fluctuations in our instrumental variables) on the wait times for new patient primary care appointments, assuming the instrumental variable assumptions are valid. The coefficients on the interactions with capacity (represented by vector $\boldsymbol{\delta}$) indicate how the capacity effect changes depending on affluence and availability of alternative health care options. If the interaction coefficients are statistically different from zero and positive (and negative for veteran unemployment), the clinician capacity effect on wait times is smaller (and demand more elastic) when more alternative health coverage options exist and veterans are potentially more affluent.

Table 2-1 provides definitions, means, and standard deviations of the variables in the analysis (Appendix Figure 6-2 shows the distributions). The average wait time across facilities in the sample is 26 days for a new patient primary care appointment. However, wait times vary substantially across VHA facilities, being over 90 days at some facilities. In terms of VHA facility capacity, there are on average 38 FTE clinicians per facility and 0.68 FTE clinicians per 1,000 enrollees. The number of private physician specialists per 1,000 capita is 2.38. Comparing the standard deviations (relative to averages), the average capacity in the VHA varies more than it does in the private sector. The average monthly nurse turnover rate across facilities is 0.0085, meaning 0.0085 nurses quit for every one nurse who is retained (among registered nurses who have less than 5 years of tenure at the VHA). Nurse turnover rates vary substantially across facility-months; the standard deviation is larger than the mean.

Table 2-1. Descriptive Statistics

Variable	Definition	Mean	SD
Wait time (days)	Mean wait time for a primary care appointment at a VHA facility in a given year-month (2^{nd} stage dependent variable)	26.43	10.78
PCP FTE	Number of full-time equivalents (FTE) of primary care physicians at a VHA facility in a given year-month	37.65	22.61
PCP FTE per 1,000 enrollees	Number of FTE primary care physicians at a VHA facility divided by the number of enrollees of that facility in 2010, multiplied by 1000 (first stage dependent variable)	0.68	0.19
Nurse turnover rate	The number of registered nurses with fewer than 5 years of tenure at a VHA facility who quit in a given year-month divided by the number of registered nurses at the beginning of the month	0.01	0.01
Non-federal specialists per 1000 capita	Number of physicians who were not primary care physicians divided by the number of people in catchment area, multiplied by 1000 (Instrumental variable #3)	2.38	0.85
% of veterans who are 65+ years old	Proportion of veterans who are 65 years old or older	0.50	0.09
% of males 18-64 years old with health insurance	Proportion of males between 18 and 64 years of age who reported having health insurance	0.80	0.07
Household median income (in $10,000s)	Median income of households in catchment area, divided by $10,000	5.31	0.99
Zillow housing price index (in $100,000s)	Zillow housing price index in catchment area, divided by $100,000	1.78	1.01
Veteran unemployment rate	Unemployment rate of veterans in catchment area	0.07	0.03
% of veterans subject to copays	Proportion of veterans who are required to pay copayments for VHA services, i.e., veterans with Priority Status 7 or 8	0.25	0.07
Medicare Advantage penetration rate	Proportion of Medicare eligible persons who are enrolled in Medicare Advantage in catchment area	0.26	0.12
Medicaid expansion indicator	Indicator that equals 1 if state had expanded Medicaid in a given year and 0 otherwise (the indicator is 0% prior to 2014 and approximately 60% by 2016)	0.23	0.42

Note: The statistics are based on our sample of 9,113 medical center-year-months – 127 medical centers spanning fiscal years 2010 to 2016.

3. RESULTS

Overall, increases in primary care clinician capacity on average (across all medical centers) were associated with lower wait times for primary care appointments (Table 3-1). The base OLS model estimates (Column [4]) indicate that a one standard deviation change in primary care clinician capacity (approximately 1 more clinician per 5,000 enrollees) is associated with lower wait times by 1.3 (=0.19 x 6.90) days. The two-stage least squares (2SLS) estimates (Columns [5] and [6]) show a larger effect of clinician capacity on wait times, which is consistent with the existence of simultaneity bias in the base OLS model. The 2SLS results in Column [5] – our preferred specification (discussion below) – indicate that a one standard deviation change in clinician capacity is associated with a 3.79 (=0.19 x 19.99) day reduction in the wait time. The 2SLS results in Column [6], which use both instrumental variables in the first stage, show similar estimates of the capacity effect.

In addition, Table 3-1 shows that the increase in R-squared by including facility fixed effects is large. Comparing Regressions [1] and [2], the R-squared increased from 0.00 to 0.48, suggesting that about half of the variation in wait times is between facilities. The main independent variable of interest – primary care clinician capacity – explains some of the remaining over-time, within-facility variation.

Table 3-1. Base Ordinary Least Squares and Two-Stage Least Squares Estimates

	[1]	[2]	[3]	[4]	[5]	[6]
	Base Model (no instruments)				IV1 Model	IV2 Model
	Coeff. S.E.	Coeff. S.E.	Coeff. S.E.	Coeff. S.E.	Coeff. S.E.	Coeff. S.E.
Constant	27.46 (0.10)***	29.15 (0.91)***	28.73 (1.01)***	33.74 (1.27)***	43.35 (4.96)***	42.84 (4.93)***
(Predicted) PCP FTE per 1000 enrollees				-6.90 (1.06)***	-19.99 (6.65)***	-19.30 (6.60)***
Medical center fixed effects		x	x	x	x	x
Year indicators			x	x	x	x
Month indicators			x	x	x	x
R-squared	0.00	0.48	0.50	0.51	0.50	0.50
Adjusted R-squared	0.00	0.48	0.50	0.50	0.50	0.50
Number of facility-year-months	9,113	9,113	9,113	9,113	9,113	9,113

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

Note: The dependent variable is the mean wait time at a VHA medical center in a given year-month. All regressions were weighted by the number of enrollees at the medical center in 2010. IV1 Model shows regression results using the instrument of the number of non-VHA specialists per 1,000 capita. IV2 Model shows regression results using two instruments – the previous mentioned instrument and the turnover rate of newly hired VHA nurses.

First Stage Results

Table 3-2 shows the regression estimates from the first stage of the instrumental variables approach. Regressions [2]-[7] include facility fixed effects, and [3]-[7] include year and month indicators, which absorb annual and seasonal national-level changes in the VHA primary care clinician capacity[16]. After including facility fixed effects, the R-squared increased from 0 to 0.81, suggesting that between-facility variation plays a major role in the variation we observe in VHA capacity, even more so than it did for wait times.

The instruments have the signs we would expect. Regressions [4]-[9] of Table 3-2 shows that the turnover rate among new VHA registered nurses is negatively correlated with primary care capacity, indicative of less appealing work environments. However, the magnitude of the relationship is small and only significant at the 5% level. A 1 percentage point increase in the nurse turnover rate is associated with a 0.0019 decrease in the number of primary care clinicians per 1,000 enrollees, which is one-hundredth of a standard deviation. Non-VHA physician specialists per capita was significantly and positively correlated with VHA clinician capacity. One more non-VHA physician specialist per 1,000 capita in the area was associated with 0.23 more VHA primary care clinicians per 1,000 enrollees. The estimate remains stable after including the control variables in our second stage model (measures of affluence, economic indicators, and availability of alternative health care coverage).

[16] Although not shown for the sake of brevity, the summer months were associated with lower capacity relative to January.

Table 3-2. First Stage of Two Stage Least Squares Estimates

	[1] Coeff. S.E.	[2] Coeff. S.E.	[3] Coeff. S.E.	[4] Coeff. S.E.	[5] Coeff. S.E.	[6] Coeff. S.E.	[7] Coeff. S.E.	[8] Coeff. S.E.
Constant	0.68 (0.00)***	0.74 (0.01)***	0.73 (0.01)***	0.23 (0.06)***	0.81 (0.10)***	0.69 (0.10)***	1.26 (0.09)***	0.81 (0.10)***
Nurse turnover rate (0-1)							-0.19 (0.09)**	-0.17 (0.08)**
Non-federal specialists per 1,000 capita				0.22 (0.02)***	0.24 (0.02)***	0.23 (0.02)***		0.24 (0.02)***
Medicaid expansion indicator					0.00 (0.00)	0.00 (0.00)	0.00 (0.00)	0.00 (0.00)
% of veterans 65+ years old (0-1)					-0.50 (0.06)***	-0.53 (0.06)***	-0.49 (0.06)***	-0.50 (0.06)***
% with health insurance (0-1)					-0.27 (0.08)***	-0.08 (0.08)	-0.28 (0.08)***	-0.26 (0.08)***
Median household income (in $10,000s)					-0.04 (0.01)***	-0.02 (0.01)**	-0.04 (0.01)***	-0.04 (0.01)***
Zillow housing price index (in $100,000s)					0.02 (0.01)***		0.02 (0.01)***	0.02 (0.01)***
Housing price index (in $100,000s)						-0.02 (0.01)**		
Veteran unemployment rate (0-1)					-0.28 (0.06)***	-0.31 (0.06)***	-0.29 (0.06)***	-0.28 (0.06)***
% of veterans subject to copays (0-1)					0.02 (0.14)	0.07 (0.14)	0.36 (0.14)***	0.02 (0.14)
Medicare Advantage penetration rate (0-1)					-0.19 (0.05)***	-0.25 (0.05)***	-0.15 (0.05)***	-0.19 (0.05)***
Medical center fixed effects		x	x	x	x	x	x	x
Year indicators			x	x	x	x	x	x
Month indicators			x	x	x	x	x	x
R^2	0.00	0.81	0.83	0.83	0.84	0.84	0.83	0.84
Adjusted R^2	0.00	0.81	0.83	0.83	0.83	0.83	0.83	0.83
Number of facility-year-months	9,113	9,113	9,113	9,113	9,113	9,113	9,113	9,113

Notes: The dependent variable is the number of full-time equivalent primary care clinicians employed by a VHA medical center in a given year-month per 1,000 enrollees. For brevity, medical center fixed effects, year, month indicators are not shown in the table.
*** p < 0.01, ** p < 0.05, * p < 0.1

We tested the strength of each instrument separately and both together. The F-statistic from using the number of physician specialists per capita alone was 97.47, much larger than the Stock and Yogo (2005) critical value for a maximum 10% 2SLS bias. The F-statistic from using nurse turnover is 0.48, suggesting a weak instrument. The F-statistic for joint significance of the two instruments (Model 9 in Table 3-2) is 50.84, which is larger than the critical value for two instruments and one endogenous variable of 19.93. These results suggest that using the number of non-VHA physician specialists per capita is a stronger instrument than nurse turnover. Moreover, including the weak instrument with the stronger one may be less preferred than simply using only the stronger instrument – the number of non-VHA physician specialists.

For this reason, we focus most of our discussion on the second stage results from using the non-VHA specialists per capita instrument, which we label as *IV1 Model*. For comparison, the tables also provide the results of the model using both instruments, which is labeled *IV2 Model*. The two models provide qualitatively similar results; however, the estimates sometimes differ in magnitude.

Demand Sensitivity to Local Area Factors

We tested whether affluence and alternative health coverage options affected the level of demand (intercept) and the sensitivity of demand (slope). In terms of level effects, the results are consistent with alternative health care options reducing the demand for VHA-provided health care. We found that Medicaid expansion, the proportion of veterans who were age-eligible for Medicare, and Medicare Advantage penetration were associated with lower VHA wait times for new patient primary care appointments. Oddly, higher proportions of males who had health insurance was associated with longer VHA wait times. This may partially be explained if veteran males who were enrolled in VHA health care also reported having health insurance; thus, the proportion may be correlated with VHA enrollment, which is indicative of VHA demand and may increase wait times.

The coefficients on the affluence measures have the expected sign, consistent with the idea that veterans who are more affluent rely less on the VHA for health care. Income and the proportion of veterans who were subject to copayments for VHA services were both negatively related with VHA wait times for primary care. Veteran unemployment was negatively related to VHA wait times. This is consistent with previous literature that documents the phenomenon that unemployed persons are less likely to seek health care treatment in general (Travers et al. 2017 and Hamad 2014). However, it is inconsistent with other studies that find that unemployment is linked to higher VHA outpatient utilization (Frakt et al. 2015), which implies that we would expect wait times to increase (if the VHA did not prepare provider supply in advance). Housing prices did not affect the level of demand in a statistically significant way in any of the models, except when the interaction term (of the housing price index and VHA capacity) was included in the model.

We found that the relationship between wait times and clinician capacity was negative, even after including the alternative health coverage and affluence measures. This means that when capacity increases (supply shifts), wait times decrease, which is consistent with a downward sloping demand curve. The estimated coefficients on the interacted terms show that the capacity effect is sensitive to three of the seven measures of affluence and alternative health care coverage (Table 3-3). The three measures that modulate the capacity effect at a given VHA facility are the proportion of veterans who are age-eligible for Medicare (Model 3), the median income in the catchment area (Model 5), and the housing price index in the catchment area (Model 6). Increases in each of these measures dampen the effect of capacity on wait times. This suggests more elastic demand when Medicare is an alternative option for veterans, income is higher, and housing prices are higher. However, when performing different combinations of these three interacted measures, we found that income was the only measure that had consistently statistically significant effects (Models 10-13).

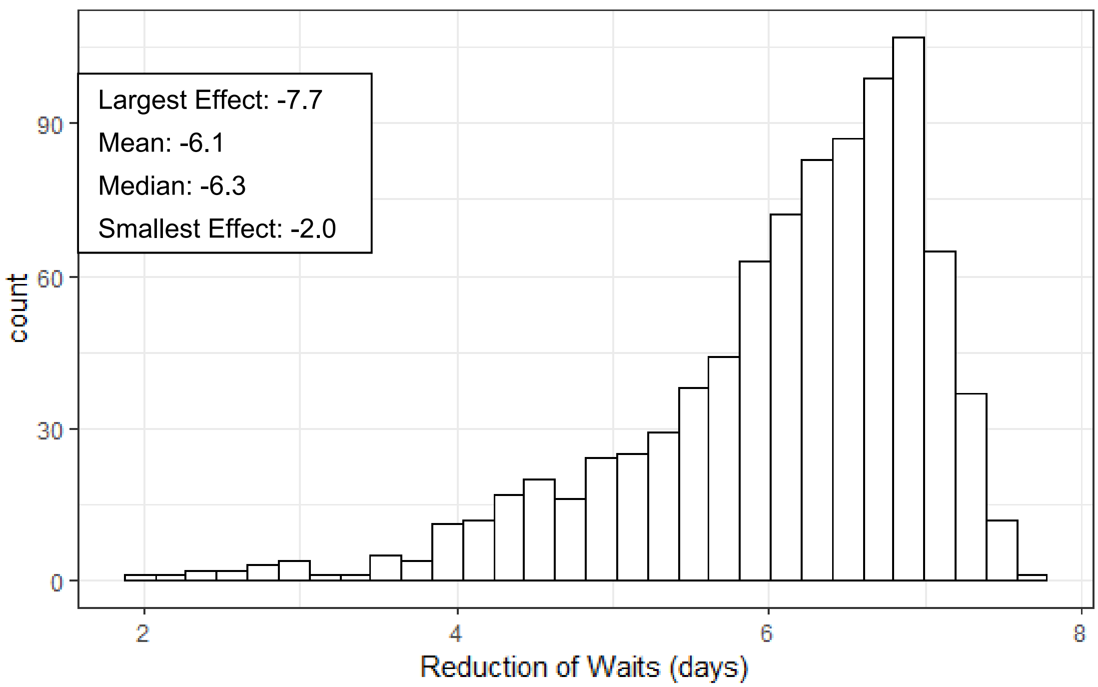
Table 3-3. Second Stage of Two Stage Least Squares Estimates

	[1] Coeff. S.E.	[2] Coeff. S.E.	[3] Coeff. S.E.	[4] Coeff. S.E.	[5] Coeff. S.E.	[6] Coeff. S.E.	[7] Coeff. S.E.	[8] Coeff. S.E.	[9] Coeff. S.E.	[10] Coeff. S.E.	[11] Coeff. S.E.	[12] Coeff. S.E.	[13] Coeff. S.E.
Constant	43.35	50.30	64.08	42.17	72.80	56.37	50.44	43.18	51.10	64.43	74.43	70.41	74.24
	(4.96)***	(14.41)***	(15.45)***	(16.87)**	(15.64)***	(14.53)***	(14.41)***	(15.09)***	(14.78)***	(15.45)***	(15.88)***	(16.19)***	(16.36)***
(Predicted) PCP FTE per 1000 enrollees	-19.99	-32.99	-49.01	-22.75	-63.40	-40.98	-32.81	-23.60	-33.99	-50.19	-64.86	-60.10	-64.39
	(6.65)***	(9.27)***	(11.32)***	(14.43)	(12.42)***	(9.60)***	(9.28)***	(11.00)**	(10.14)***	(11.33)***	(12.66)***	(13.67)***	(13.84)***
… x % 65+ years old			30.09							19.77	8.54		21.12
			(12.22)**							(12.90)	(14.45)		(15.48)
… x % with health insurance				-11.14									
				(12.03)									
… x Household median income					5.47						4.91	4.47	5.25
					(1.49)***						(1.76)***	(2.28)**	(2.69)*
… x Zillow housing price Index						4.04				3.38		1.13	0.99
						(1.28)***				(1.35)**		(1.96)	(1.99)
… x Veteran unemployment rate							-9.39						
							(20.85)						
… x % Veterans with Priority 7 or 8								-46.46					
								(29.32)					
… x Medicare Advantage penetration rate									3.06				-33.92
									(12.60)				(15.60)**
Medicaid expansion indicator		-1.08	-1.02	-1.11	-0.97	-0.97	-1.07	-1.10	-1.07	-0.95	-0.97	-0.96	-0.97
		(0.33)***	(0.34)***	(0.34)***	(0.34)***	(0.34)***	(0.34)***	(0.34)***	(0.34)***	(0.34)***	(0.34)***	(0.34)***	(0.34)***
% of veterans 65+ years old (0-1)		-1.74	-23.73	-0.51	-6.18	-5.66	-2.21	-3.12	-1.85	-19.46	-11.97	-6.47	-21.18
		(7.33)	(11.55)**	(7.45)	(7.42)	(7.43)	(7.40)	(7.38)	(7.34)	(11.68)*	(12.29)	(7.44)	(12.97)
% with health insurance (0-1)		55.48	54.68	63.59	54.24	58.35	55.43	55.86	55.46	57.35	54.14	55.27	54.71
		(8.06)***	(8.06)***	(11.90)***	(8.06)***	(8.11)***	(8.06)***	(8.06)***	(8.06)***	(8.13)***	(8.06)***	(8.26)***	(8.26)***
Median household income (in $10,000s)		-6.56	-6.71	-6.49	-10.10	-6.73	-6.59	-6.69	-6.58	-6.80	-9.79	-9.51	-9.88
		(1.00)***	(1.00)***	(1.00)***	(1.39)***	(1.00)***	(1.00)***	(1.00)***	(1.00)***	(1.00)***	(1.49)***	(1.73)***	(1.91)***
Zillow housing price index (in $100,000s)		0.57	0.43	0.64	0.16	-2.66	0.54	0.53	0.58	-2.23	0.16	-0.67	-0.84
		(0.59)	(0.59)	(0.59)	(0.60)	(1.18)**	(0.59)	(0.59)	(0.59)	(1.21)*	(0.60)	(1.56)	(1.57)
Veteran unemployment rate (0-1)		-23.70	-23.13	-23.75	-22.37	-23.50	-16.90	-23.24	-23.55	-23.16	-22.35	-22.56	-23.60
		(7.39)***	(7.39)***	(7.39)***	(7.39)***	(7.38)***	(16.82)	(7.39)***	(7.41)***	(7.39)***	(7.39)***	(7.40)***	(7.41)***
% of veterans subject to copays (0-1)		-21.29	-23.05	-22.77	-15.63	-18.44	-20.54	17.19	-21.14	-20.07	-16.70	-15.86	-18.05
		(13.29)	(13.30)*	(13.38)*	(13.37)	(13.31)	(13.39)	(27.68)	(13.30)	(13.35)	(13.49)	(13.37)	(13.51)
Medicare Advantage penetration rate (0-1)		-24.92	-26.27	-24.26	-27.28	-25.98	-25.23	-25.60	-27.00	-26.69	-27.43	-27.15	-5.29
		(4.91)***	(4.94)***	(4.96)***	(4.95)***	(4.92)***	(4.95)***	(4.93)***	(9.89)***	(4.94)***	(4.95)***	(4.95)***	(11.27)
R-squared	0.50	0.51	0.51	0.51	0.51	0.51	0.51	0.51	0.51	0.51	0.51	0.51	0.51
Adjusted R-squared	0.50	0.50	0.50	0.50	0.50	0.50	0.50	0.50	0.50	0.50	0.5	0.5	0.5
Number of facility-year-months	9,113	9,113	9,113	9,113	9,113	9,113	9,113	9,113	9,113	9,113	9,113	9,113	9,113

Notes: The dependent variable is the mean wait time at a VHA medical center in a given year-month. All regressions include medical center fixed effects, year indicators, and month indicators. All regressions were weighted by the number of enrollees at the facility in 2010. For brevity, medical center fixed effects, year indicators, and month indicators are not shown in the table.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

To illustrate the range of variation in the estimated capacity effect across areas, Figure 3-1 is a histogram of the capacity effect (reduction in wait time) due to a one standard deviation increase in capacity. The effect is based on the estimates in Model 5 and the observed values of median local income levels[17]. The capacity effect ranges from 2.0 to 7.7 days, meaning that a one standard deviation (0.19) increase in capacity (or approximately 1 FTE per 5,000 enrollees) would yield wait time reductions between 2.0 and 7.7 days, depending on the income in the surrounding area of a VHA facility. The reduction is lower in places with higher income, and it is highest in areas with low income. This suggests that the demand for VHA care is more inelastic (with respect to wait times) in less affluent areas.

Figure 3-1. Histogram of Capacity Effect, Given Observed Values of Income



Note: This is a histogram of the magnitude of the reduction in wait time for a one standard deviation increase in capacity using Model 5 and the observed values of income.

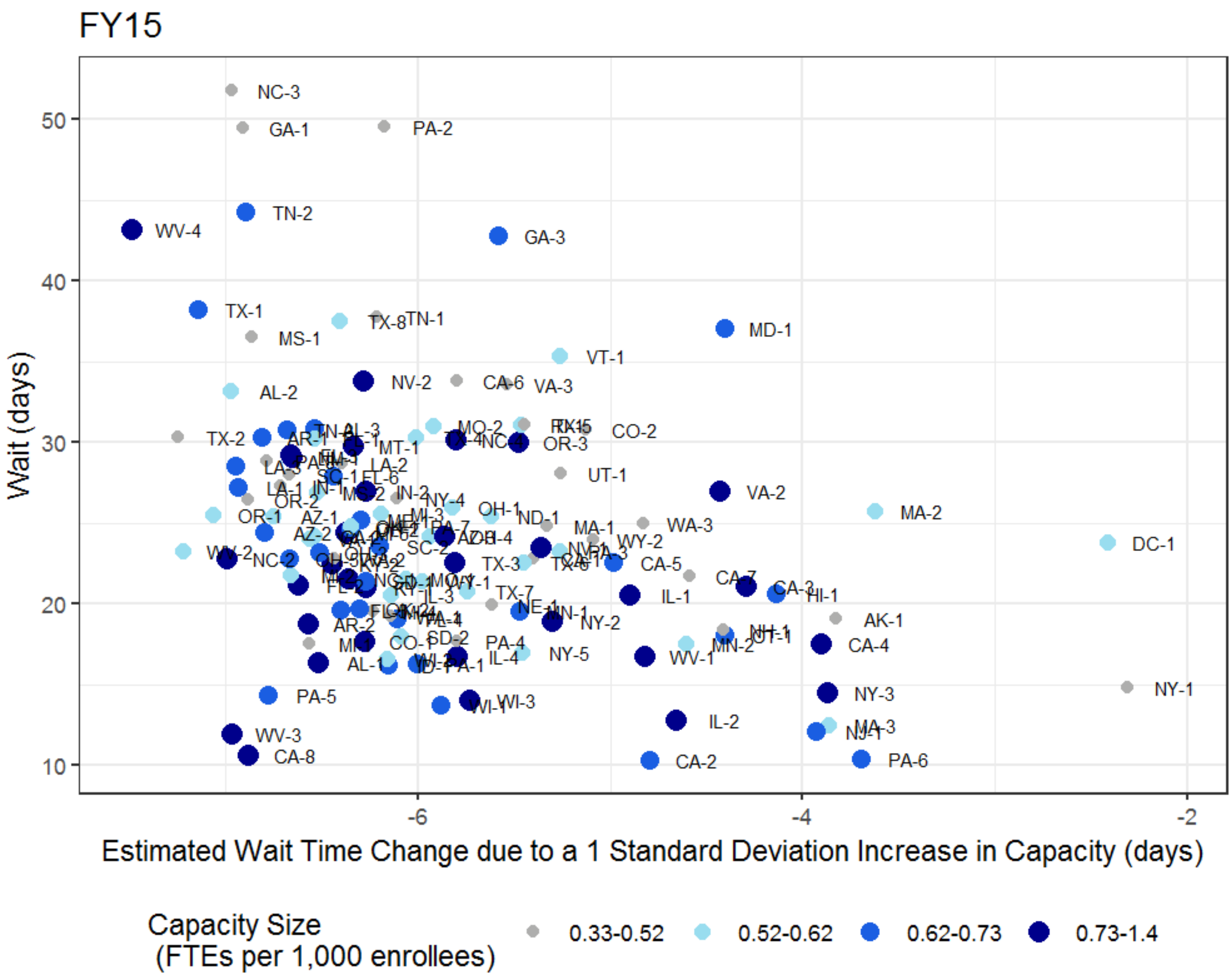
[17] The variation in income is what drives the variation in the capacity effect, which can be seen by the similar shape of the distribution to the Income distribution in the Appendix.

Simulation Exercise

The VHA has a fixed budget, so the model could be used to target resources. If the goal is to reduce wait times, the VHA could most efficiently reduce them by hiring clinicians at facilities that not only have high wait times but also are highly responsive to additional clinicians – in other words, facilities that have large capacity effects.

Figure 3-2 plots facilities according to their wait time and capacity effect. The facilities in the upper left quadrant have higher wait times (e.g., >30 days on y-axis) and are more responsive to capacity (e.g., >7.6 day capacity effect on the x-axis, given a one standard deviation increase in capacity). The size and color of the dot indicate the capacity size in 2015, divided into quartiles. The base capacity size helps us understand the relative size of a one standard deviation increase in capacity (or 0.19 FTEs per 1,000 enrollees). For example, a VHA facility in North Carolina (NC-3 in upper left quadrant) has a capacity size of less than 0.52 primary care FTEs per 1,000 enrollees; so a 0.19 increase would be around a 50% increase in capacity. A facility in West Virginia (WV-4 in upper left quadrant) has one of the largest capacity sizes – greater than 0.73 – so a 0.19 increase in capacity would not be proportionally as large. In the Appendix, we look at what it would take to achieve a 30-day maximum wait time across all facilities.

Figure 3-2. Wait Times, Estimated Capacity Effects, and Capacity Sizes of VHA Facilities (FY2015)



4. DISCUSSION

Our results indicate that a one standard deviation increase in capacity (approximately 1 clinician FTE per 5,000 veteran enrollees or 28% of the sample mean) leads to a 3.79-day reduction in wait time for primary care (35% of the standard deviation (10.78 days) and 14.3% of the sample mean facility-month wait time (26.43 days)). The effect of capacity on wait times varies geographically due to variations in income and, depending on the specification, Medicare availability and housing prices. The more affluent an area, the smaller the capacity effect and the more elastic the demand. When wait times increase in these areas, veterans seek care from providers outside the VHA.

These results are consistent with previous findings that show the interdependency between VHA health care and other public health care coverage options. If veterans have access to coverage from Medicaid, Traditional Medicare, or Medicare Advantage, they use VHA health care less (Hebert et al. 2018, Liu et al. 2018, Frakt et al. 2015, Pizer and Prentice 2011). We find that having access to these payers is associated with lower wait times for primary care.

This is particularly relevant for implementation of the Veterans Choice Act and MISSION Act, which emphasize access to care and direct VHA funds to pay for private health care for veterans. It may be most effective for the VHA to focus its funds and provide more care in areas with large capacity effects. In areas with small capacity effects, the VHA may be better off relying on community providers. Currently, the VHA is using a version of the basic model presented in this paper (without interactions) to comply with the MISSION Act Title IV Section 401. In this Section, the VHA must identify underserved areas and develop a strategic plan to improve access. With our model, the VHA is identifying areas in which the explained component of wait times is high. In these areas, the wait time is high due to (the lack of) VHA clinician supply, availability of private health coverage, affluence, economic factors, and/or other market factors that may affect the wait time for VHA health care.

Comparison to Previous Demand Elasticity Estimates

While the capacity effect is useful for policymakers, it is not directly comparable to previous studies that estimate demand elasticities. Translating the capacity effect to a demand elasticity is highly sensitive to functional form assumptions. Instead of making these assumptions, we compute the elasticity of appointments with respect to wait times at the average wait time and number of appointments: $\frac{\partial A}{\partial C} / \frac{\partial W}{\partial C} \times \frac{W}{A}$. To get $\frac{\partial A}{\partial C}$, we estimated an OLS and 2SLS regression of appointments on capacity. We included the control variables for local affluence, economy, and alternative health coverage options (no interacted terms), facility fixed effects, year indicators, and month indicators. Using no instrument for capacity (OLS), we found the elasticity to be -0.52 (223.4 / -33.0 * 26.4 / 356.7). Using 2SLS (similar to Model 2 in Table 3-3), the elasticity was -1.03 (458.6 / -33.0 * 26.4 / 356.7).

This is larger than the -0.32 elasticity found by Windmeijer, Gravelle, and Hoonhout (2004), who studied First General Practitioner Outpatient Visits in Scotland using data from the year 2000. This is the only study that estimated the demand elasticity for primary care appointments; the rest were on surgical procedures or hospital admissions. They showed that a 1% increase in lagged wait time by 2 months led to a 0.32% decrease in the number of new outpatient visits per patient in a practice. See the Appendix for a detailed comparison of our estimate and that of previous studies.

Our estimate may be larger than this previous estimate for several reasons. First, we mitigated the downward simultaneity bias using instruments. Second, we study the U.S. VHA, which differs from Scotland's delivery system in important ways, such as: the larger size of the private health care market in the U.S. and the type of patients and health conditions treated. Third, we use a slightly different quantity measure – number of new patient appointments rather than new patient visits per patient in the practice.

5. CONCLUSION

This study develops new theoretical and empirical approaches to the relationship between the supply of health care providers and access to care. The results indicate that a shift in supply can yield different predictions for wait times depending on the elasticity of demand. It demonstrates that a resource-constrained health care system, such as the VHA, could improve wait times without increasing its budget by re-allocating clinical resources from markets with highly elastic demand to markets with less elastic demand, which have a higher yield in terms of improving access. The differential benefit from additional public investment is due to the VHA operating in markets with a variety of other public and private alternatives and serving an economically heterogeneous population. The findings have direct implications for current VHA policy in determining where needs are greatest and how the VHA should respond to those needs.

6. APPENDIX

Veterans Health Administration Background

In 1930, President Hoover created the Veterans Administration (VA) to administer all services to veterans, including health care, pensions, employment opportunities, and other services. The program became known as the Department of Veterans Affairs in 1988. Today, there are 21.6 million U.S. veterans (National Center of Veteran Analysis and Statistics, 2015). Many reside in the southern part of the U.S. (U.S. Census Bureau, 2010). They often have more health issues and a higher incidence of chronic illnesses (e.g. diabetes, hypertension, cancer) than the civilian population (Kramarow and Pastor, 2012).

The Veterans Health Administration (VHA) is one of three divisions within the Department of Veterans Affairs. It is a network of 168 federally-owned and operated hospitals, 1,053 community-based outpatient clinics, and 135 nursing homes (Department of Veteran Affairs "About VA", 2016). The proposed 2020 budget for the VHA is $79.1 billion. In 2014, Congress passed the Veterans Access, Choice, and Accountability Act (or Choice Act), which appropriated $10 billion to be used between 2014 and 2017 to improve the access to non-VHA providers (Congressional Budget Office, 2014; Department of Veteran Affairs "Budget in Brief", 2017). Since 2014, the budget for the VHA has increased approximately $5 billion per year, in effort to bolster VHA staffing and infrastructure and improve access to VHA care. The MISSION Act of 2018 extends funding to the Choice program of $5.2 billion for a year as the community (non-VHA) care program is set up (Department of Veterans Affairs "Budget in Brief", 2018). Using this budget, the VHA pays for the salaries of clinicians and staff, maintains VHA-owned clinics, and sometimes builds new clinics.

Of the 21.6 million U.S. veterans, 9.1 million are enrolled in VHA health benefits, and approximately 5.9 million use VHA medical care each year (Bagalman, 2014). Although only 27% of veterans receive VHA care, the number of veterans who use VHA health care has grown 51% since 2001 (Bagalman, 2014). To be eligible for VHA health benefits, a veteran now needs to have served for 24 continuous months or be discharged due to a military service-connected disability. Unlike Medicare and other types of health care coverage, the VHA does not require veterans to pay a premium to receive benefits. Furthermore, copayments are relatively small for veterans, depending on their disability status and need for financial assistance (Table 6-1 and Table 6-2).

Table 6-1. VHA Priority Group Classification of Veterans

Priority Group	Description
1	Service-connected disability rating of 50% or more; unemployable due to service-connected disability
2	Service-connected disability rating of 30% or 40%
3	Former Prisoners of War, Purple Heart or Medal of Honor recipients, service-connected disability rating of 10% or 20%, discharged for a disability from being in the line of duty
4	Catastrophically disabled; receiving aid or assistance from VA
5	Not disabled; income below VA threshold; receiving VA pension benefits; eligible for Medicaid
6	Not disabled; served in specific wars
7	All else whose income is below VA threshold and who agree to pay copayments
8	All else whose income is above VA threshold and who agree to pay copayments

Notes: The lower the Priority Group number the more benefits a veteran receives. The 2016 income thresholds range from $31,978 to $44,968 for veterans who have 0-4 dependents. The threshold increases by $2,198 for each additional dependent. Source: Department of Veteran Affairs "VA Benefits: Priority Groups" (2016).

Table 6-2. Premiums and Cost-sharing By Public Health Care Program

	Priority Group 1	Priority Group 2-6	Priority Group 7	Priority Group 8	Medicare	Medicaid
Annual premium	$0	$0	$0	$0	Part B: ~$1260 (depends on income) Part A: ~$0	Typically $240-$360
Deductible	$0	$0	$0	$0	Part B: $166 Part A: $1288 per episode	
Primary care visit	$15	$15	$15	$15	Part B: 20% coinsurance	$4 to 20% coinsurance
Specialist visit	$50	$50	$50	$50	Part B: 20% coinsurance	$4 to 20% coinsurance
Inpatient services	$0	$0	$256 for 1st 90 days, $128 for each additional 90 days + $2/day	$1288 for 1st 90 days, $644 for each additional 90 days + $10/day	$0 for 1st 60 days, $322/day between 61 and 90 days, $644/day after 90 days	$75 per episode to 20% coinsurance
Medication	$0	$8 per 30-day supply with $960 annual cap*	$9 per 30-day supply with no annual cap*	$9 per 30-day supply with no annual cap*	Depends on private plan	$4 to 20% coinsurance

*for non-service-connected conditions

Source: Department of Veteran Affairs "VA Benefits: Health Benefit Copays" (2016).

Even though the VHA provides less expensive health care coverage, many eligible veterans seek care from community, non-VHA providers. There are several reasons for this phenomenon. First, VHA hospitals and clinics may not be as conveniently located as community providers. Second, the choice of VHA providers and treatments may be more limited than alternative coverage options. Third, VHA appointments may have longer wait times than the alternatives and so treatment may be delayed. Veterans must balance low financial cost against distance, choice, and wait times, relative to his or her alternatives.

Wait Times

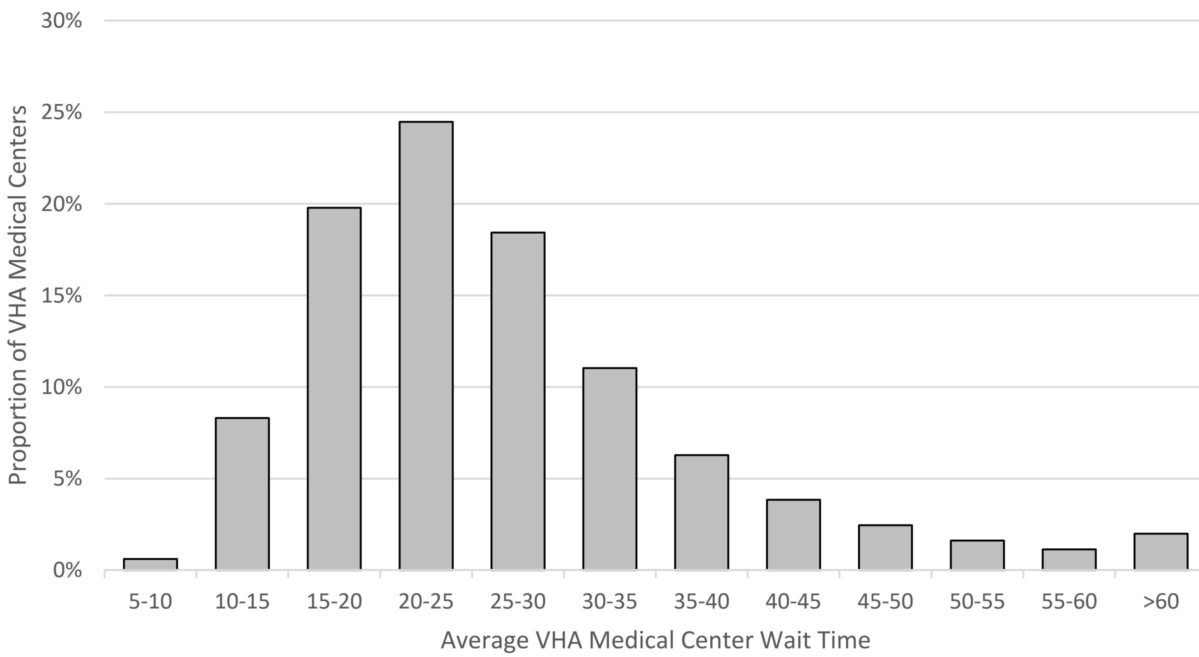
The media has highlighted the long wait times for veterans to see a clinician, for example, the VHA scandal of 2014 (VA Office of Inspector General, 2014). Providing minimal costs for patients under a finite budget, which limits the number of providers, a public system like the VHA often needs to ration health care through wait times (Lindsay and Feigenbaum, 1984). In order to maximize population health under this constraint, sometimes it is prudent to prioritize certain patients (or areas) over other patients (Gravelle and Siciliani, 2009).

To address the wait time issue, policymakers have set maximums for wait times. In 1995, the VHA set a maximum of 30 days and required most of its medical centers to meet that goal within three years. A 2001 government report showed that VHA medical centers were having problems meeting the 30-day goal (U.S. Government Accountability Office, 2001). In 2007, the VHA was sued for "failing to provide adequate and timely benefits and medical care," according to Veterans for Common Sense, an advocacy group for veterans (2014). In 2011, the wait-time requirement was reduced to 14 days for many types of appointments, despite the already present difficulties in meeting the 30-day requirement and the future increase in the number of veterans who would be returning from Afghanistan and Iraq (Brunker, 2014).

There have also been issues in the record-keeping of wait times. The VHA previously measured wait times by the days between the 'desired date' for an appointment and the date of that appointment. In some facilities (supposedly under pressure to meet the 30- and 14-day wait times), schedulers misrecorded the desired date to match the date of the appointment, which would make the wait time zero. In this study, we address this problem by not using the desired date to measure wait times, but rather using the date that a new appointment was entered into the system. This latter date, called the ‘create date’ cannot be manipulated by schedulers.

Our data indicate that the average wait time for a new patient primary care appointment in 2014 was approximately 22 days. Wait times vary substantially across VHA medical centers (Figure 6-1)

Figure 6-1. Histogram of Wait Times Among VHA Medical Centers 2014



Additional Figures and Tables

Table 6-3. Source of Variables

Variable Used in Analysis	Source	Source Level of Aggregation
Average wait time	VHA CDW	Clinic Visit
Number of enrollees	VHA Enrollee Table	Enrollee-year
% of veterans 65 years old and older	VHA Enrollee Table	Enrollee-year
% of veterans with Priority Status 7 or 8	VHA Enrollee Table	Enrollee-year
Clinician FTEs	VHA OPES	VAMC-month
Low-tenured registered nurse turnover rate	VHA Human Resources	VAMC-month
Number of non-VHA physician specialists (all physicians minus PCP)	AHRF	County-year
Median household income	AHRF	County-year
% of males 18-64 years old with health insurance	AHRF	County-year
Population	AHRF	County-year
Zillow Home Value Index	Zillow	County-month
Veteran unemployment rate	ACS	County-year
Medicare Advantage penetration rate	CMS	County-month
Medicaid expansion indicator	Kaiser Family Foundation	State-year

VHA: Veterans Health Administration. CDW: Corporate Data Warehouse. OPES: Office of Productivity, Efficiency, and Staffing. VAMC: Veterans Affairs medical center. AHRF: Area Health Resource File. ACS: American Community Survey. CMS: Centers of Medicare and Medicaid Services.

Figure 6-2. Distribution of variables in model

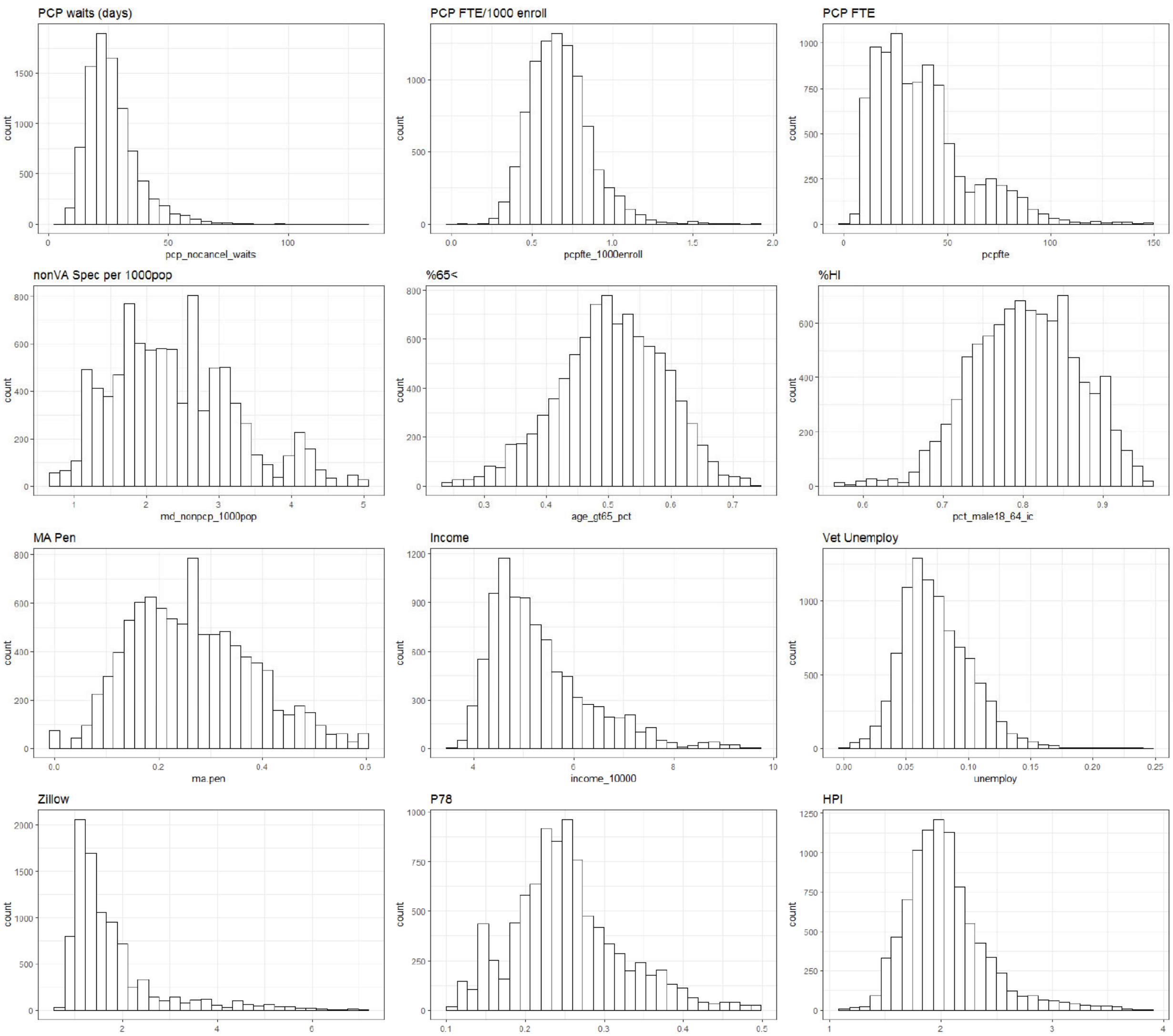
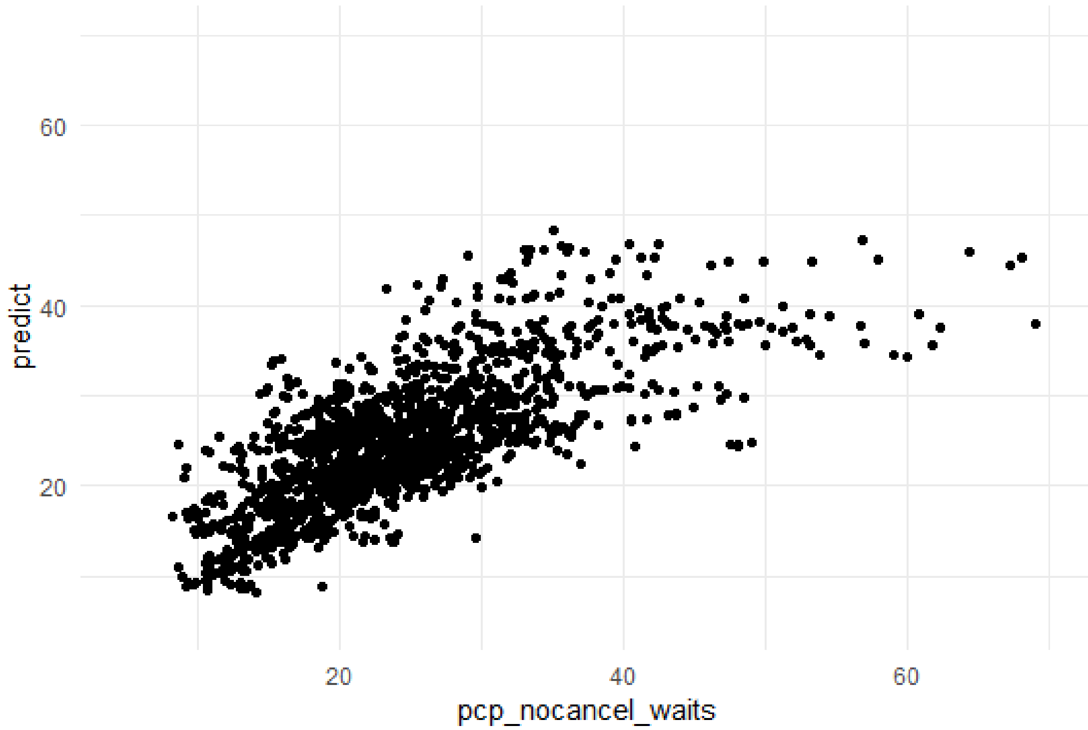


Figure 6-3. Predicted vs. Actual Wait Times (Regression [6]) FY 2014

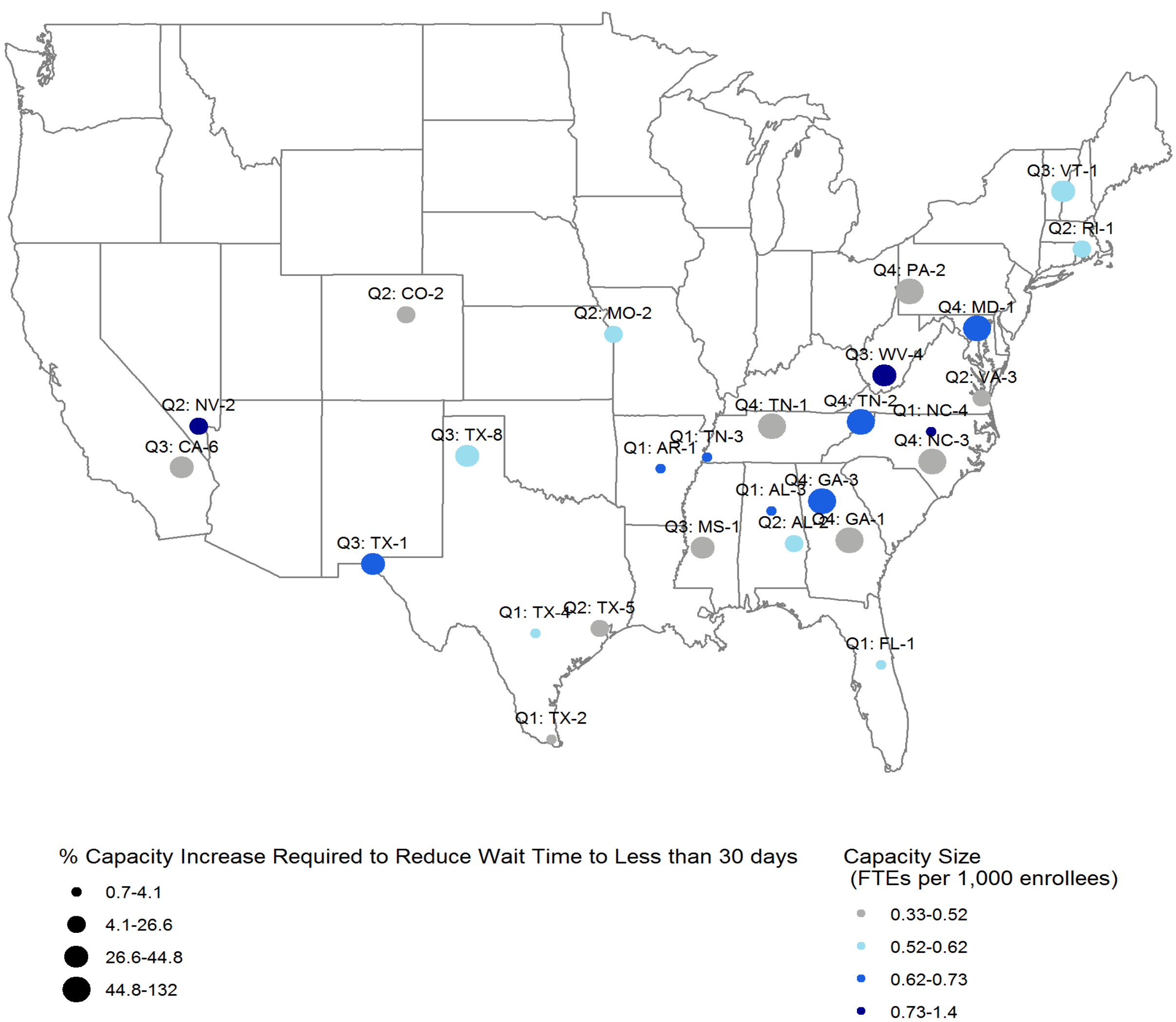


Additional Simulation Exercise

The map (Figure 6-4) illustrates a simulation exercise: how much capacity is needed to lower average wait times to the current VHA mandated target of 30 days? The map shows facilities with greater than 30-day wait times. The dot size represents the percent increase in capacity that would be needed in order for the facilities' average wait time to decrease to 30 days. The color represents the base capacity in 2016. The percent increase is dependent on the facility's sensitivity to capacity changes (or the capacity effect), baseline average wait time, and baseline capacity. For example, the facility NC-3 (Fayetteville, NC) had the highest average wait time for new patient primary care appointments in 2015 (Figure 3-2). It also had one of the lowest primary care clinician capacities. In order to bring its average wait time down to 30 days, the facility would need to increase its capacity by 132 percent, putting it into the fourth quartile in terms of percent increases in capacity. To illustrate demand elasticity differences, facilities TX-1 and MD-1 have similar capacity sizes (0.62-0.73 primary care clinicians per 1,000 enrollees) and similar wait times (around 37 days from Figure 3-2). However, MD-1 is less sensitive to a change in capacity (more elastic demand). Thus, it would be easier to lower the wait times in TX-1 than in MD-1, and efficient allocation would suggest that TX-1 should have priority over MD-1 when allocating the budget for hiring primary care clinicians.

Interestingly, the map shows that many of the facilities that have greater than 30-day average wait times are located on the southeast part of the U.S. Some of this may be attributed to the low capacity relative to the number of veterans enrolled in the VHA (gray dot). Some might be attributed to high wait times, e.g. TN-2 and GA-3 (Figure 3-2). However, there are several facilities for which the wait time is not that much higher than 30 days and is simply not that sensitive to capacity increases (demand is more elastic), e.g. MD-1.

Figure 6-4. Map of Facilities with >30-day Waits and Capacity Increase to Reduce Their Waits to 30 Days.



Comparison to Previous Literature

Recall from the Discussion Section, our estimate of the demand elasticity of appointments with respect to wait times is -0.52 or -1.03, depending on whether capacity was instrumented in the regression of appointments on capacity. Table 6-4 shows the previous literature's estimates in reverse chronological order. Our estimate is larger than previous estimates, including Windmeijer, Gravelle, Hoonhout (2004) that was highlighted in the Discussion Section. The table provides both the demand elasticity with respect to wait times and, whenever possible, estimates of the effect of supply variables on wait times (be they based on a reduced form model or on the first stage model for a structural demand equation). Although some elasticity estimates are not directly comparable to the way we define measures in this study (e.g., output quantity and supply measure), five of the eight are in the range of 0.15% to 0.35% reduction in utilization for a 1% increase in wait time. The lowest is 0.09 and the highest 1.5. Thus, our estimate is within a similar range, however, larger than the bulk of estimates. Moreover, the evidence suggests that the elasticity varies geographically. Using the range of capacity effects (-11.34 to -43.65), the elasticity (using OLS) would range from -0.38 to -1.45 or (using 2SLS) from -0.77 to -2.99.

Table 6-4. Previous Literature's Estimates of Demand Elasticity and Capacity Effect

Study	Medical service	Country & Data Years	Demand Elasticity	Capacity Effect
Riganti, Sicilliani, Fiorio (2017) Table 2 & A1	9 elective surgical procedures	Italy, 2010-2014	1% increase in wait leads to a 0.15% decrease in discharges per 1000 capita	1% change in beds leads to a 0.00576% change in wait
Sicilliani and Martin (2007)	Elective surgeries	England, 1999-2001	n/a	Increase in competition by one hospital (from 5 to 6 at sample mean) reduces wait times by 0.5 weeks (17 to 16.5 weeks)
Windmeijer, Gravelle, Hoonhout (2004) Table 4	First GP outpatient visits	Scotland, 2000	1% increase in wait time lagged 2 months leads to a 0.3168% decrease in first outpatient visits 'rate' (i.e., per patient in practice)	n/a
Gravelle, Smith, and Xavier (2003), Table 2 & page 97	Elective surgery	England, 1987-1993	1% increase in proportion of people waiting >3 months lagged by 1 quarter leads to a 0.21 to 0.30% decrease in additions to the wait list	n/a
Martin and Smith (2003), Table 5	Elective surgery in 7 specialties	England, 1991-1998	Actual waits being 1 ppt above expected wait leads to 0.09 ppt lower actual utilization relative to expected utilization	Study discusses estimating this effect but does not show it (i.e., first stage)
Gravelle et al. (2003)	Admissions to acute care hospitals	England, 2000-2001	1% increase in wait leads to a 1.5% decrease in admissions*	n/a
Gravelle, Dusheiko, and Sutton (2002), page 18	Cataract procedures	England, 1995-1998	1% increase in wait leads to 0.25% decrease in odds ratio in the admissions rate per 10,000 capita**	n/a
Martin and Smith (1999)	Admissions to hospitals for elective surgery	England, 1991-1992	Actual waits being 1 ppt above expected wait leads to 0.21 ppt lower utilization relative to expected utilization	Computed: 1% increase in beds leads to a 0.242% decrease in wait times

*Gravelle et al. (2003) do not estimate the elasticity directly. However, we computed it for comparison to other papers. Gravelle et al. (2003) found that 1 more day in wait time leads to 0.017 fewer admissions (Table 2 page 996); the mean wait was 91 days and mean admissions was 1.0; thus, a 1% increase in wait (i.e., 0.91 day) leads to a 0.017 x 0.91 / 1.0 = 0.015 decrease in admissions.

** Gravelle, Dusheiko, and Sutton (2002) show estimated coefficients from a regression of ln(Admission rate / 1-Admission rate) on the median wait time with imputed values for 25% of the practices. They separately report the demand elasticity; however, the interpretation of the outcome variable is unclear, whether it is the admission rate or the odds ratio of the admissions rate.

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