More Doctors in Town Now? Evidence from Medicaid Expansions

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Abstract

We examine how physicians' practice locations are affected by Medicaid expansions. We focus on the dramatic Medicaid eligibility expansions for pregnant women that took place between the early 1980s and the early 1990s. Instrumenting for eligibility using a simulated measure of Medicaid generosity, we estimate that the increase in the number of pregnant women eligible for Medicaid between 1982 and 1992 raised the supply of obstetricians and gynecologists (OB/GYNs) by 14 percent in poor counties. Our results are driven by recently graduated OB/GYNs and are concentrated in counties with urban populations. Our results demonstrate that Medicaid coverage rules are an important determinant of physician supply.

JEL Codes: I18, I11, I13, I38, H51, J22, J44
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1 Introduction

Socioeconomic disparities in the U.S. are an important factor for the adverse maternal and infant health outcomes. For example, the high U.S. infant mortality is attributable to high postneonatal mortality, which is largely due to poor birth outcomes among individuals of low socioeconomic status (Chen et al., 2016). In an effort to improve health outcomes of this group of population, Medicaid—a federal-state matching entitlement program—covers medical expenses for people with limited income and resources, in which two-thirds of the program enrollees are from low-income families (Gruber, 2000). Between the early 1980s and the early 1990s, states implemented dramatic Medicaid eligibility expansions for pregnant women, which extended coverage to include individuals in a much broader range of low-income families.

However, the benefits of these expansions will be limited if beneficiaries have difficulties in gaining access to care. According to the Centers for Medicare & Medicaid Services, Medicaid spending accounted for 16 percent of the total national health expenditures in 2019 (CMS, 2019). Despite substantial resources put into the Medicaid program, improvement in access to and quality of health care for its beneficiaries has often been inadequate.[1] An important reason for the slow-improving outcomes is a constraint on physician labor supply. Offering insurance coverage and financial support does not ensure access to health care services, especially if physicians are not available to care for eligible individuals (Fossett and Peterson, 1989; Fossett et al., 1992; Fingar et al., 2015). In fact, the shortage and maldistribution of physicians have been major concerns regarding physician supply (Currie and Thomas, 1995; Ku et al., 2011; Bodenheimer and Smith, 2013). Expanding public health insurance programs leads to a larger demand for health care, and physicians may have financial incentives to capture the pool of new patients. However, physician capacity constraints often limit physicians' ability and availability to provide services to meet the demand. Moreover, policymakers may worry about the potentially negative "spill-overs" on existing insured patients due to the expansions, such as longer wait times, lower quality of care, and increased physician burnouts (Garthwaite, 2012; Buchmueller et al., 2016). Thus, it is important to

[1] A large body of literature has been devoted to assessing the efficacy of the Medicaid program. The findings are generally mixed, but somewhat positive. Buchmueller et al. (2015) provide a comprehensive summary of research on the impact of Medicaid on various outcomes.

understand the impact of changes in public health insurance coverage on physicians' labor supply decisions. Existing research on physician behavior in response to Medicaid expansions generally examines the effects on the *intensive* margin or changes in the amount of care offered to an individual patient and provider participation in the Medicaid program (Baker and Royalty, 2000; Garthwaite, 2012; Buchmueller et al., 2016).[2] Yet, there is limited scope for physicians to adjust on the intensive margin, and the total growth in the supply will also need to come from the *extensive* margin or an increase in the number of working physicians.[3] While both margins clearly play critical roles in determining access to services and shaping public health, the latter has been much understudied.

In this paper, we study how obstetricians and gynecologists (OB/GYNs) make their practice location decisions in response to Medicaid eligibility expansions for pregnant women. Specifically, we examine whether and how the stock of county-level OB/GYNs changes as eligibility rates for pregnant women vary over time during 1982–1992. We exploit the fact that these expansions were implemented at very different rates across states during the period and base our analysis on a national sample of county-level physician counts from the American Medical Association (AMA). Following Currie and Gruber (1996), we use a simulated eligibility rate—that employs variation in Medicaid eligibility due to changes in Medicaid coverage rules—as an instrument for the actual eligibility rate in the fixed effects model. It is plausibly exogenous, as the simulated fraction is only based on the differences in Medicaid eligibility rules across states and is independent of other underlying characteristics of states. We find that an increase of 27 percentage points in eligibility—the average change in the actual fraction eligible over our sample period—resulted in about a 14 percent increase in OB/GYNs per capita in poor counties. The effects are strongly associated with urbanicity of counties. Moreover, the resulting growth in OB/GYNs is mainly driven by an increased supply of recently graduated physicians in these areas. Generally, we estimate larger effects when net Medicaid incentives are greater.

[2] One notable exception is a recent study by Huh (2021), which shows the impact of Medicaid expansions on the number of providers. Huh (2021) focuses on adult Medicaid dental expansions and examines dentist supply as the outcome of interest in a difference-in-differences-in-differences framework.

[3] Changes in the number of physicians may come from physicians joining existing practices or hospitals/clinics, or starting new practices. In theory, public health insurance could crowd out private insurance that is more lucrative (Cutler and Gruber, 1996). It is therefore unclear ex ante how Medicaid expansions may affect the supply of physicians on the extensive margin.

We provide the first empirical evidence for the impact of Medicaid eligibility expansions in the 1980s and 1990s on the extensive margin of physician supply. We add to the recent study by Huh (2021) that examines how adult Medicaid dental expansions affected practice locations of dentists. Our findings show that large-scale Medicaid expansions targeted at a specific population can affect location decisions of the physicians who are most likely to provide health care services to that specific group. More generally, these results point to a possible role of Medicaid in influencing the supply and geographic distribution of physicians. This has major implications for policymakers, especially when facing various issues associated with the current shortage and maldistribution of physicians in the U.S.

The rest of the paper proceeds as follows. Section 2 discusses the institutional background and relation of our paper to the literature. Section 3 describes the main datasets and empirical approach. Section 4 presents results and Section 5 concludes.

2 Background

2.1 Medical policy changes

Prior to the 1980s, Medicaid eligibility for pregnant women was closely linked to participation in the Aid to Families with Dependent Children (AFDC) program, which only covered a group of narrowly defined families subject to stringent income and resource requirements. The AFDC benefits were generally available to single-parent families, and the income limits varied by state and were often very low. For instance, in 1979, the income eligibility level for a family of four in Texas was equivalent to 24 percent of the federal poverty line (FPL) (Currie and Gruber, 1996).

Starting in the early 1980s, states began extending coverage to a much wider set of families by weakening the linkage to AFDC eligibility. For example, the income threshold for eligibility was no longer based on the AFDC limit but on the FPL instead, and the requirements on family structure were removed. Moreover, states were granted with discretion to the income cutoffs for determining eligibility. In 1987, the average income cutoff across states was approximately 60 percent of the FPL (Cutler and Gruber, 1996). After a series of expansions, by the early 1990s, all states had offered Medicaid coverage to pregnant women with family incomes below 133 percent of the FPL, regardless of family composition.[4] On average, eligibility rates for pregnant women increased from about 15 percent to 42 percent between the early 1980s and the early 1990s (see Figure 1). The expansions took place at much different rates across states during the time period. These differences provide substantial variations in eligibility across states and over time, constituting key sources for our identification.[5]

2.2 Relation to literature

This paper complements the literature by studying the effect of public health insurance expansions on physician behavior. A vast literature focuses on the demand-side outcomes, such as health care access, utilization, and quality, whereas research on the impact on health care providers—the supply side of the market—has been relatively limited (Buchmueller et al., 2015).[6] Eligibility expansions can affect the extent to which and the way physicians offer medical services. Baker and Royalty (2000) study the effect of Medicaid expansions for pregnant women in the late 1980s and find that the increased eligibility improved access to health care only from public physicians, those working in hospitals and public clinics. Garthwaite (2012) investigates the impact of the State Children's Health Insurance Program and finds that more pediatricians reported accepting and seeing Medicaid patients, but with a decrease in the average time spent on patients. Buchmueller et al. (2016) examine how dentists respond to expanding adult Medicaid dental coverage. They show that the expansions increased participation of dentists in the Medicaid program and provision of Medicaid dental services without substantially reducing time spent on patients, which was achieved mainly by making greater use of dental hygienists.

All of these studies shed light on whether and how physicians adapt to the new or expanding insurance programs on the *intensive* margin. However, the supply of medical services by providers

[4] Currie and Gruber (1996) and Buchmueller et al. (2015) provide more details on the evolving expansions during this time period.

[5] The concurrent expansions also increased coverage to children, teenagers, or other family members. Our main dataset, discussed in Section 3.1, includes detailed information only for OB/GYNs and not for other medical specialties. In this paper, thus, we focus on pregnant women, who were the main target of the policy changes during the time period and experienced the largest growth in eligibility rates.

[6] Buchmueller et al. (2015) review how Medicaid has evolved in recent decades and summarize related studies that examine the impact of Medicaid on a wide range of outcomes, including both health and non-health related outcomes.

depends heavily on the *extensive* margin—the number of working physicians—as well. A recent paper by Huh (2021) studies the impact of adult Medicaid dental expansions on dentist supply and finds that the expansions increased the number of dentists per capita in poor counties by 13 percent. On the other hand, our paper examines whether physicians enter or locate to areas in response to one of the largest, yet narrowly targeted Medicaid expansions. These expansions resulted in a substantial demand shock for health care among pregnant women, the targeted population (Currie and Gruber, 1996, 2001). For example, (Currie and Gruber, 1996) show that the probability of delayed prenatal care among women eligible for Medicaid was reduced by almost 50 percentage points. As with dentists, in theory, physicians may respond to the expansions on the extensive margin by joining existing practices or hospitals/clinics (i.e., with more job posts to hire more physicians in response to the increased demand), or starting new practices.[7] We provide important policy implications for potentially shaping the geographic distribution of health care services, which relies on location decisions of physicians.

Our paper is also related to a host of literature that studies physicians' decisions on where to practice.[8] A series of recent papers have developed models to characterize the choice of location by physicians. They provide evidence that physicians are generally responsive to financial incentives, such as student loan forgiveness programs and physician subsidies for Health Professional Shortage Areas, when choosing practice locations (Zhou, 2017; Falcettoni, 2018; Kulka and McWeeny, 2018; Khoury et al., 2020). We examine the role of Medicaid incentives in location decisions of physicians, whose major patients include the targeted population by the expansions, and we also explore the underlying mechanisms.[9] Results of our analysis shed light on an important aspect of physician decision-making on labor supply.

[7]The latter option is more costly, especially given high overhead costs ranging from 40 to 60 percent of total practice revenues (Woolhandler and Himmelstein, 1991; Kletke et al., 1996). Due to data unavailability, we are unable to distinguish between physicians who joined existing practices and those who started new practices during our study period. In 1983, the fraction of physicians in owner practices of 10 or less was nearly 80 percent, and the fraction of solo-owner physicians was 44 percent (Khullar et al., 2018).

[8]See Bärnighausen and Bloom (2009) for a comprehensive review of the literature on financial-incentive programs and their impact on practice location decisions among health care providers.

[9]In 1987, about a quarter of all patients treated by physicians (private and public combined) were poor patients on Medicaid (Baker and Royalty, 2000).

3 Data and Empirical Strategy

3.1 Data

We obtain data on OB/GYNs from the Physician Masterfile established by the American Medical Association (AMA). The Masterfile contains profile records for all physicians in the U.S., including those who have completed or are in the process of completing requirements to practice medicine. Our main data contain county-level counts of OB/GYNs from this database for the years 1982–1992, including those working/employed in privately-owned practices, public clinics, or hospitals.[10] We also have information on graduation year from medical school to investigate heterogeneity in physician response by career stage.

Data on the actual and simulated eligibility measures for Medicaid come from Currie and Gruber (1996). Eligibility rates are defined as the fraction of female residents aged between 15 and 44 who are eligible for Medicaid at the state level. We use the simulated fraction of this variable as an instrument in our instrumental variables (IV) analysis (see Section 3.2). The simulated eligibility rate is constructed by applying Medicaid eligibility rules in each state and year to a fixed national sample from the Current Population Surveys (CPS).[11] We also obtain Medicaid-to-private fee ratios for obstetrical care from Currie et al. (1995). Eligibility rates/fee ratios are not available for the following five states: Alaska, Arizona, Kentucky, Texas, and Wyoming.

We supplement our data using the Area Health Resources Files (AHRF) and the Surveillance, Epidemiology, and End Results (SEER) Program, which provide demographic variables at the county level, including per-capita income, unemployment rate, fraction of female population aged 15–44, fraction of black population, and female (ages 15–44) population density. These variables follow Huh (2021) and are important indicators for local conditions that may simultaneously affect the structure of the Medicaid program and providers' decisions on practice locations. In addition, we obtain county-level poverty rates in 1980 and 1990 from the U.S. Census Bureau, and the 1983 urban-rural indicators from the Economic Research Service (ERS) at the U.S. Department of

[10] The years 1984 and 1990 are excluded due to data unavailability. We do not have information to distinguish between private and public OB/GYNs, but results in our paper are presumably driven by private OB/GYNs. For instance, in 1987, nearly 90 percent of all OB/GYNs were private physicians (Baker and Royalty, 2000).

[11] Currie and Gruber (1996) provide more details on how to construct this measure.

Agriculture (USDA).

Figure 1 shows aggregate trends in the actual fraction of pregnant women aged 15–44 who are eligible for Medicaid between 1982 and 1992.[12] The fraction goes up from about 15 to 42 percent, nearly a three-fold increase during this time. The figure also illustrates an increasing pattern in the supply of OB/GYNs over time, changing from about 22 to 28 OB/GYNs per 100,000 females aged 15–44.

In Figure 2, we plot these trends in OB/GYN supply separately by quartiles of states with varying expansion sizes as measured by the percent increase in eligibility rates from 1982 to 1992. Per-capita OB/GYN supply rises from 21.16 to 25.52, or by 20.6 percent, among the first quartile (i.e., smallest expansion) of states. On the other hand, there is a 36.7 percent respective increase from 32.41 to 44.29 among the fourth quartile (i.e., largest expansion) of states. The second and third quartiles of states experienced a 17 percent increase and a 27 percent increase in OB/GYN supply, respectively.

Table 1 presents the summary statistics for the main variables used in our analysis. We show the overall average for all counties, and also separately for metropolitan and non-metropolitan counties.[13] The average number of OB/GYNs is 26.31 per 100,000 women aged 15–44 per county. Our paper also examines whether the impact on location decisions differs between new physicians—who recently graduated and completed medical residency—and those with more experience and established practices.[14] We categorize a physician's stage of career according to the estimated number of years in practice: at most 5 years defined as early-career (i.e., capturing newly graduated physicians), between 6 to 20 years as mid-career, and those beyond as late-career physicians. On average, mid-career OB/GYNs account for nearly 50 percent of the total supply, followed by late-career OB/GYNs, making up about 28 percent. We observe more OB/GYNs per capita in metropolitan counties than in non-metropolitan counties, and the distribution of physician supply across career stages is similar between these two groups of counties. The remaining rows of Table 1 report the statistics for county-level controls. Non-metropolitan areas tend to have lower income,

[12] Appendix Table A1 reports these rates across years.

[13] We follow the rural-urban continuum codes adopted by the USDA's ERS. See https://www.ers.usda.gov/topics/rural-economy-population/rural-classifications/what-is-rural.aspx.

[14] Currently, residencies for OB/GYNs usually last four years in length. This was also true during our study period.

higher unemployment, smaller fractions of blacks and females aged 15–44 in the population.

3.2 Empirical strategy

Our estimation exploits the variations in eligibility rates due to Medicaid expansions that occurred in a staggered fashion between the early 1980s and the early 1990s. The dramatically expanded coverage to pregnant women resulted in a greater demand for prenatal care and delivery services. We examine whether and the extent to which these eligibility changes may have affected physicians' decisions on where to practice. We regress county-level OB/GYN counts per 100,000 on a Medicaid eligibility index—the actual fraction of pregnant women aged 15–44 who are eligible for Medicaid in their state of residence by year. Specifically, our estimating equation is as follows:

$$Y_{ct} = \beta_0 + \beta_1 ER_{s(t-1)} + \gamma X_{ct} + \theta_c + \delta_t + \varepsilon_{ct}, \tag{1}$$

where Y_{ct} denotes the number of OB/GYNs per capita in county c in year t; $ER_{s(t-1)}$ stands for the lagged eligibility rate in state s in year t; X_{ct} includes a set of county-level characteristics as discussed above; county fixed effects, θ_c, control for any county-invariant factors; and year fixed effects, δ_t, control for omitted variables, such as new procedures or treatments for women's health, that develop over time mostly at the national level.[15] The key variable of interest is eligibility rates, and its coefficient β_1 measures the average change in per-capita OB/GYN counts across all counties in a state, given a unit change in the fraction eligible at the state level. We use a lagged measure of this variable to allow time for physicians to respond to the expansions. A positive coefficient estimate would imply a net increase in the supply of OB/GYNs as a result of the expansions, who could work in/establish private practices to capture the new customers or accept hiring offers in hospitals and public clinics that face an increasing demand for related services.

Concerns arise regarding the endogeneity of our key variable of interest. As pointed out by

[15] All models also include a control for lagged Medicaid-to-private fee ratios for obstetrical care, but our estimated effects of the eligibility changes without this variable are materially the same. It is an important factor affecting physician behavior but likely to be endogenously determined. An increase in Medicaid-to-private fee ratios may result from a relative increase in Medicaid fees, but this may also indicate a relative decrease or a relatively small increase in private fees. We still include it as a control variable, as our goal is to estimate the causal relationship between the eligibility changes and physician supply. Currie and Gruber (1996) also consider the fee ratio variable in their estimation and find little difference in results with and without it.

Currie and Gruber (1996) and Garthwaite (2012), unobserved economic and demographic characteristics of the local environment may simultaneously affect the scale and scope of the Medicaid program as well as physicians' practice location decisions. Ignoring these sources of variations may induce a spurious correlation between the outcome variable and eligibility measure. To address this endogeneity problem, we use the simulated fraction eligible as an instrument for the actual eligibility rate, as done in a number of prior studies (Currie and Gruber, 1996; Gruber and Simon, 2008; Miller and Wherry, 2019). Specifically, based on a fixed national sample of 3,000 women from the CPS, the fraction eligible of this sample is constructed according to each state's eligibility rules by year. By doing so, the calculated fraction only depends on the variations in the legislative environment and is independent of other underlying characteristics of states. Taken together, our estimation strategy relies on a fixed effects approach combined with an IV technique, so as to minimize the potential bias in the estimated effects. Our first-stage results show a strong correlation between the endogenous variable and the instrument.[16] Each observation is a county-year combination, weighted by the size of female population aged 15–44 in the estimation. We cluster standard errors at the state level to allow for dependence in the residuals across counties within a state (Drukker, 2003).

As our preferred specification, we estimate heterogeneous effects of the eligibility changes based on different types of counties. This helps better understand the underlying mechanisms for OB/GYNs' response to Medicaid expansions. The unbalanced within-state distribution of physicians across counties shows that urban or metropolitan areas are generally preferred to rural or non-metropolitan areas for practice locations. Urbanicity and income are in fact among the major factors in physicians' location decisions (Hurley, 1991). As a result, even with the same level of Medicaid expansions, physicians' response may vary across counties within a state. To take into account this important factor in the decision-making process, we examine whether physicians' location choice in response to the expanded coverage varies between metropolitan and non-

[16] Coefficient estimate on the simulated eligibility rate, using Equation (1) with the actual eligibility rate as the outcome, is 0.8726 and it is highly significant (p-value < 0.001). The partial F-statistic is large (2,156), which indicates a strong instrument. First-stage results for our IV estimation are available in Appendix Table A2.

metropolitan areas. Specifically, we estimate the following equation:

$$Y_{ct} = \beta_0 + \beta_1 ER_{s(t-1)} + \beta_2 ER_{s(t-1)} \times Metro_c + \gamma X_{ct} + \theta_c + \delta_t + \varepsilon_{ct}, \tag{2}$$

where $Metro_c$ is equal to 1 if county c is classified as metropolitan and 0 otherwise according to the 1983 ERS rural-urban continuum codes. The coefficient β_1 measures the baseline effect of the expanded coverage on physician supply for non-metropolitan areas, and the sum of the coefficients β_1 and β_2 measures the corresponding effect for metropolitan areas. The constant effect of the metropolitan status is absorbed by county fixed effects. Other variables are defined as in Equation (1).[17]

Moreover, we examine the effect by county poverty rates. By construction, high-poverty counties are more likely to see a higher take-up rate, compared to low-poverty counties.[18] If physicians respond to the eligibility expansions, we would expect a greater inflow of OB/GYNs toward poorer counties in the expansion states with greater Medicaid incentives. Therefore, we estimate the effects separately for counties with varying poverty rates. Importantly, this allows the differential effects between metropolitan and non-metropolitan counties to also differ by the poverty level.

In addition to the geographic variations, physicians' own characteristics, career stage in particular, tend to affect their response to the expansions. For example, recent graduates are generally more likely to consider Medicaid incentives in their location decisions and thus have greater incentives to enter or locate to places with an increased demand due to the expansions. On the other hand, physicians who have settled in their current location or established practices would be less responsive to demand shocks, because the opportunity cost of relocating is relatively high. State-specific licensure requirements add to the cost of moving across states. These differential effects by career stage are frequently observed in the literature (Polsky et al., 2000; Kessler et al., 2005; Khoury et al., 2020; Huh, 2021). Thus, we run the equations separately for early-career, mid-career, and late-career OB/GYNs.

[17] The first-stage partial F-statistic using Equation (2) is also large (1,099).

[18] The rate of take-up measures the rate of enrollment in Medicaid among individuals who are newly eligible in the expansion state.

4 Results

4.1 Main results

Table 2 reports coefficient estimates for the key variables from Equations (1) and (2) for all OB/GYNs. The coefficients of interest are β_1 on the fraction eligible, and $\beta_1 + \beta_2$ on the sum of the fraction eligible and its interaction with the metropolitan indicator. OLS estimates are shown in Columns (1) and (2), and IV estimates in Columns (3) and (4). Overall, OLS and IV estimates are quite similar. In the first panel of the table, we consider all counties regardless of poverty and find that the average effect of the fraction eligible on overall OB/GYN supply using Equation (1) is positive but insignificant (Columns (1) and (3) for OLS and IV estimates, respectively).[19] This effect is averaged across all counties within states and thus does not adequately capture differential within-state responses to the eligibility changes. Results from our preferred specification, Equation (2), are reported in Columns (2) and (4). These results take into account differential effects of the expansions on location decisions by the metropolitan status. Estimates in the first panel suggest that the impact on non-metropolitan counties is negative while it is positive for metropolitan counties of the expansion states, but both are insignificant.

Next, we extend our main analysis by examining differential effects by local poverty levels. Specifically, we group counties into tertiles of poverty rates based on the 1980 and 1990 Censuses: low poverty, moderate poverty, and high poverty.[20] A pooled model is then estimated where every control variable is interacted with the poverty indicator. Results are reported in the second to fourth panels of Table 2. The average poverty rate for counties with low poverty is 9.16 percent. This increases to 14.24 percent and 23.58 percent for counties with moderate poverty and high poverty, respectively. First using Equation (1), estimates reported in Columns (1) and (3) show that the effect of the expansions is concentrated in counties with moderate poverty. For example, the IV estimation in Column (3) suggests an increase of about 2.81 ($= 0.1045 \times 26.85$) OB/GYNs per

[19]Each column in the first panel of Table 2 corresponds to a separate regression, all with the same number of observations using the full sample.

[20]In the appendix, we show that results are robust when using alternative categories of poverty rates, such as quartiles or quintiles. We use a categorical (ordinal) measure of poverty rates for the following reasons: to avoid reliance on the linear relationship between physician supply and poverty rates; to be more robust to measurement errors in the poverty rate variable; and to have easier interpretations of results.

100,000 in these counties in response to a 27-percentage-point increase in eligibility, or by 13.64 percent based on the mean of 20.60 physicians per 100,000.[21] The impact on counties with either low or high poverty is imprecise and small in magnitude. Moreover, we find using Equation (2) that among counties with moderate poverty, the increase in OB/GYN supply is largely driven by metropolitan areas. Indeed, if there exists a positive relationship between physician supply and Medicaid expansions, the increase of the supply would most likely occur in metropolitan areas, which is consistent with the observed urban concentration among health care professionals. This is shown in Columns (2) and (4) in the third panel. Our IV estimation in Column (4) suggests that Medicaid expansions increased the supply in metropolitan counties with moderate poverty by 3.54 ($= 0.1320 \times 26.85$) OB/GYNs per 100,000, or by 10.38 percent based on the mean of 34.11 physicians per 100,000. OLS estimates in Column (2) are slightly smaller in magnitude, when compared to IV estimates. We focus on results from the IV estimation below, although we report OLS estimates in all tables for comparison.

The remainder of Column (4) in Table 2 shows that the effect of the expansions on non-metropolitan counties is insignificant across different poverty levels, except for counties with low poverty, which is marginally significant at the 10 percent level and is negative. This finding might suggest an interesting pattern of potential physician flows in response to the changes in eligibility rules: a within-state movement away from non-metropolitan counties and possibly towards metropolitan counties with a high take-up rate. In other words, we may observe responses not only across states, but also across counties where physicians are increasingly choosing metropolitan areas instead of non-metropolitan areas of the expansion states. Evidence of such physician flows across counties is particularly relevant as we are estimating the impact on the supply in the short run. We are not estimating general equilibrium effects at the national level, which is in line with Malani and Reif (2015) and Huh (2021).

Assuming that the increased demand for obstetric care is driven by women who actually enroll in Medicaid, we can discuss the following local average treatment effect (LATE) of Medicaid expansions on physician supply: the ratio of the intent-to-treat and first-stage effects. About 30 percent of women who gained eligibility during our sample period enrolled in Medicaid (Currie

[21] The average change in the actual fraction eligible over the sample period, 1982–1992, was 26.85.

and Gruber, 1996). Thus, an estimate of LATE would be about a 35 percent ($= 10.38 \div 0.30$) increase in the supply. More precisely, this estimate needs to be adjusted downward (upward) if the actual take-up rate is higher (lower) than 30 percent in metropolitan counties with moderate poverty.

We show that simply focusing on the overall impact across all counties may underestimate the value of Medicaid expansions. Again, if physicians respond to the expansions on the extensive margin, we would expect an increase of physician counts in poor and densely populated areas, where the demand shock is large enough with a high take-up rate. At the same time, our results from Table 2 suggest that physicians might be unwilling to locate to areas with poverty that is too high. However, we cannot rule out the possibility that metropolitan counties with high poverty experienced an increase in OB/GYN supply as much as counties with moderate poverty did. We illustrate this comparison in Panel (a) of Figure 3, where we plot coefficient estimates for all metropolitan counties and for metropolitan counties with different poverty levels along with the corresponding 95-percent confidence intervals. The confidence interval for counties with high poverty is wide and includes the estimate of 0.1320 for counties with moderate poverty. In contrast, the confidence interval for counties with low poverty rules out this point estimate.

Table 3 presents results using Equation (2) for all counties and by poverty for early-career, mid-career, and late-career OB/GYNs.[22] The estimated effect for all OB/GYNs from Table 2 is mostly driven by early-career OB/GYNs. These early-career physicians are the ones who recently completed residency with at most five years of experience in practice. IV estimates in Column (4) suggest that the number of early-career OB/GYNs per 100,000 in metropolitan counties with moderate poverty increased by 1.72 ($= 0.0640 \times 26.85$), or by 19.44 percent based on the mean of 8.85 physicians per 100,000. Again, the positive effect on physician supply is concentrated in metropolitan counties with moderate poverty, and the difference by poverty is statistically distinguishable from zero only in comparison to counties with low poverty. This is shown in Panel (b) of Figure 3. For non-metropolitan areas with low or high poverty, estimates are negative and insignificant or marginally significant. On the other hand, results in the first panel of Ta-

[22] A complete set of results where we replicate Table 2 for OB/GYNs in each career stage is reported in Appendix Tables A3–A5.

ble 3 for all non-metropolitan counties show that early-career OB/GYN supply decreased by 1.08 ($= -0.0403 \times 26.85$), or by 17.45 percent based on the mean of 6.19 physicians per 100,000. While speculative, this might suggest that the negative effect on physician supply for non-metropolitan areas of the expansion states is spread out across varying poverty rates.[23]

Columns (5) and (6) of Table 3 report IV estimates for mid- and late-career OB/GYNs, respectively. The estimated coefficients for these two groups of physicians are insignificant and imprecise. The only exception is the effect for metropolitan counties with moderate poverty among late-career physicians, which is marginally significant at the 10 percent level. Estimates suggest an increase in the supply of physicians by 11.70 percent in these counties of the expansion states, equivalent to 1.14 ($= 0.0424 \times 26.85$) late-career OB/GYNs per 100,000. This is possibly due to delayed retirement.[24] A visual representation of largely insignificant results for non-early-career OB/GYNs is shown in Panels (c) and (d) of Figure 3.

In all, our findings suggest that the resulting changes in the stock of OB/GYNs are mainly driven by those newly graduated. This is generally aligned with the literature. Physicians at the early stage of career, especially recent graduates, are more likely to consider location decisions in response to external market forces, such as expansions of public health insurance programs (Huh, 2021), insurance market penetration (Polsky et al., 2000), changes in malpractice laws or policies (Kessler et al., 2005), and financial and salary incentives (Khoury et al., 2020). In contrast, later-career physicians face a higher cost of relocating in the presence of external variations and have less incentive to move.

4.2 Additional heterogeneity results

Focusing on early-career OB/GYNs, we next estimate differential effects of the eligibility expansions by the size of existing supply of OB/GYNs at the county level. Specifically, we divide counties into those with an above-median supply of OB/GYNs and those with a below-median supply of

[23] Appendix Figure A1 plots coefficient estimates for non-metropolitan counties for all OB/GYNs and for OB/GYNs by career stage. The effects are generally indistinguishable across different poverty levels.

[24] This pattern of decision-making among late-career providers is also found in previous studies (Kessler et al., 2005; Huh, 2021).

OB/GYNs in the beginning of our sample period.[25] We estimate our preferred specification, Equation (2), separately for counties with low (below-median) and high (above-median) existing supply of OB/GYNs. We use existing supply as a proxy for the extent of available practices already in place that could hire new OB/GYNs in response to the expansions.[26] Table 4 reports the resulting estimates by poverty for these two types of counties. The effects are concentrated among counties with an above-median supply of existing OB/GYNs as shown in Column (4), where the estimated increase in the supply in metropolitan areas with moderate poverty is 2.05 ($= 0.0765 \times 26.85$) OB/GYNs per 100,000. This corresponds to a 19.23 percent increase relative to the mean of 10.66 physicians per 100,000. However, the difference between these counties with low versus high existing supply of OB/GYNs is statistically insignificant. So, we cannot rule out the possibility that low-supply areas experienced an increase in physician supply as much as high-supply areas did.

As discussed previously, physicians may respond to the expansions by joining existing practices or hospitals/clinics, or by starting their own practices. However, the latter option is often more costly, especially given the high overhead costs ranging from 40 to 60 percent of total practice revenues (Woolhandler and Himmelstein, 1991; Kletke et al., 1996). The expansions likely generated greater Medicaid incentives to respond among those with lower overhead costs. While our data do not allow us to distinguish between existing and new practices, our findings in Table 4 are consistent with our prediction that more new OB/GYNs could have responded to Medicaid expansions by joining existing practices or hospitals/clinics.[27]

We now estimate heterogeneous effects of the expansions for counties in states with above-median and below-median Medicaid-to-private fee ratios.[28] Again, we acknowledge the potential endogeneity associated with these fee ratios. Higher fee ratios do not necessarily indicate better practice environments in terms of payments for OB/GYNs, especially because our data include mostly private physicians. At the same time, however, OB/GYNs who actually responded to the increase in eligibility rates should have preferred high Medicaid reimbursement rates. While high

[25]For existing supply, we only count mid- and late-career OB/GYNs.

[26]We still assume that physicians are the final decision-makers in terms of where to go and start a new practice or join an existing practice to maximize their earnings.

[27]Of note, this also could be picking up some of the unobserved differences in practice environments between counties.

[28]The national average fee ratios changed from about 55 percent to 58 percent during our sample period.

fee ratios do not always indicate high Medicaid reimbursement rates, they likely generated greater Medicaid incentives for OB/GYNs on average in response to the expansions, when compared to low fee ratios. If fee ratios were very low, then the impact of the eligibility expansions on physician supply would be minimal, if anything. This is because low fee ratios are due to either low Medicaid reimbursement rates or high private reimbursement rates, both of which should not drive the effects of the eligibility expansions we have estimated. Results presented in Table 5 test this implication. We find that the effects are concentrated in states with high (above-median) fee ratios, where we estimate an increase in the supply by 26.32 percent, or 2.39 ($= 0.0889 \times 26.85$) OB/GYNs per 100,000 for metropolitan counties with moderate poverty (see Column (4)). Again, the differences in comparison to states with low fee ratios are not statistically significant, but they still provide suggestive evidence that is consistent with how physicians may be expected to react to Medicaid expansions, given the level of fee ratios in the expanding states.[29]

4.3 Robustness checks

In the appendix, we perform a set of robustness exercises for our main results for early-career OB/GYNs by poverty using Equation (2). Appendix Table A7 reports estimates without controlling for any county-specific characteristics and fee ratios to show that our results are not sensitive to the choice of control variables or exclusion of fee ratios. We also provide estimates from unweighted regressions in Appendix Table A8 and using non-lagged eligibility rates in Appendix Table A9. These estimates are all materially the same as our main results. In addition, Appendix Table A10 presents estimates using alternative categories of poverty (i.e., quartiles and quintiles). Results based on quartiles and quintiles of poverty are consistent with the main conclusion and suggest that the effects are largely concentrated in areas with potentially a large enough demand shock due to Medicaid expansions (i.e., metropolitan counties with relatively high poverty).

[29]In the appendix, we also examine heterogeneity across states with respect to the presence of tort reforms. Many of these laws implemented in the 1980s and 1990s reduce medical liability for physicians in practice and therefore, if anything, have a positive impact on the supply of physicians. So, it is possible that Medicaid expansions could have been more effective in attracting OB/GYNs if the expanding states also had implemented tort reform laws. Results shown in Appendix Table A6 suggest a larger and more precise impact for states with tort reforms.

5 Conclusion

We study how physicians consider location decisions in response to the dramatic expansions of Medicaid eligibility to pregnant women during the years 1982–1992. The impact of Medicaid expansions on the extensive margin has been understudied in the literature, despite its potentially important policy implications. Overall, OB/GYN counts per capita increased by about 14 percent in poor counties after the expansions. We find that the observed changes in physician supply are driven by early-career OB/GYNs or recent graduates and are concentrated in metropolitan counties. Interestingly, within an expansion state, the resulting changes may partially arise from an inflow of new OB/GYNs from non-metropolitan counties to metropolitan counties with high take-up rates.

These findings suggest that Medicaid expansions indeed have the potential to improve access to care for the targeted population, not just by increasing coverage, but also by altering the geographic distribution of physicians. Our results are consistent with the predictions and findings in prior studies (Buchmueller et al., 2015; Huh, 2021). At the same time, however, they also imply that the effects of the expansions are not equal across counties and career stages. This heterogeneity is important from a policy perspective. To better address the issue of physician capacity constraints or shortages in specific areas of the country, a more thoughtful program would need to take into account the varying incentives and costs of different providers in labor supply and location decisions.

It has been a long-standing national priority to improve maternal and infant health outcomes, but the U.S. still falls behind other developed countries along these dimensions (Jacob, 2016; Ranji et al., 2019). While we study the effect of a policy change that took place decades ago, its policy implications remain important. Medicaid expansions we consider entail one of the largest but narrowly targeted expansions of public health insurance programs thus far. There is a lack of existing research examining the impact of these expansions on the extensive margin of supply. Ultimately, our findings show a positive and strong correlation between physician supply and the eligibility changes, and we add important empirical evidence to the current literature that evaluates Medicaid.

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Table 1: Summary statistics for key variables

Variable	All counties Mean	All counties Std. Dev.	Metro counties Mean	Metro counties Std. Dev.	Non-metro counties Mean	Non-metro counties Std. Dev.
OB/GYNs per 100,000	26.31	139.54	32.38	24.33	24.34	159.93
Early-career OB/GYNs per 100,000	6.68	38.37	8.19	8.38	6.19	43.88
Mid-career OB/GYNs per 100,000	12.18	66.90	14.60	12.01	11.39	76.67
Late-career OB/GYNs per 100,000	7.45	39.94	9.58	8.73	6.76	45.67
Income per capita (in $)	12574.94	3737.72	14706.45	4339.35	11883.74	3231.80
Unemployment rate (in %)	7.66	3.49	6.57	2.61	8.01	3.66
Population density (per square mile)	55.56	440.35	199.44	874.20	8.91	10.31
Population female aged 15-44 (in %)	21.58	2.56	23.81	1.98	20.85	2.30
Population black (in %)	11.20	15.66	11.84	12.07	10.99	16.66
N	23,592		5,777		17,815	

Notes: This table reports unweighted means for the main outcome and control variables, across all years 1982–1992 excluding 1984 and 1990 due to data unavailability. The metropolitan status follows the 1983 rural-urban continuum codes.

Table 2: Effect of eligibility changes on OB/GYN supply by poverty, all OB/GYNs

	Outcome: OB/GYNs per 100,000			
	OLS estimation		IV estimation	
Coefficient	**(1)**	**(2)**	**(3)**	**(4)**
All				
Fraction eligible	0.0179	-0.0359	0.0178	-0.0365
	(0.0303)	(0.0330)	(0.0295)	(0.0319)
Fraction eligible		0.0521		0.0517
+ Fraction eligible × Metropolitan		(0.0445)		(0.0422)
Low poverty				
Fraction eligible	-0.0037	-0.0720	-0.0109	-0.0802*
	(0.0436)	(0.0446)	(0.0421)	(0.0436)
Fraction eligible		0.0080		0.0020
+ Fraction eligible × Metropolitan		(0.0462)		(0.0448)
Moderate poverty				
Fraction eligible	0.1013***	0.0050	0.1045***	0.0086
	(0.0317)	(0.0366)	(0.0302)	(0.0349)
Fraction eligible		0.1262***		0.1320***
+ Fraction eligible × Metropolitan		(0.0363)		(0.0341)
High poverty				
Fraction eligible	-0.0439	-0.0781	-0.0342	-0.0688
	(0.1055)	(0.1114)	(0.0986)	(0.1035)
Fraction eligible		0.0331		0.0388
+ Fraction eligible × Metropolitan		(0.1346)		(0.1281)
County fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
County controls	Yes	Yes	Yes	Yes
Observations	23,592	23,592	23,592	23,592

Notes: This table reports OLS and IV estimates of Equations (1) and (2) for all OB/GYNs by tertiles of county poverty rates. The average poverty rates in 1980 and 1990 are used. The coefficients of interest are β_1 on the fraction eligible, and $\beta_1 + \beta_2$ on the sum of the fraction eligible and its interaction with the metropolitan indicator (for Equation 2). Other regressors include county-specific characteristics (i.e., income per capita, unemployment rate, population density, fraction of 15-44-year-old female population, and fraction of black population), fee ratios, county fixed effects, and year fixed effects. Each column in the first panel corresponds to a separate regression. For each column in the remaining panels, a pooled model is estimated where every regressor is interacted with the poverty indicator. Standard errors in parentheses are clustered at the state level. Estimates are weighted by the county's female population, ages 15–44.

* $p < 0.10$ ** $p < 0.05$ *** $p < 0.01$

Table 3: Effect of eligibility changes on OB/GYN supply by poverty, OB/GYNs by career stage

	Outcome: OB/GYNs per 100,000					
	OLS estimation			IV estimation		
	Early career	Mid-career	Late career	Early career	Mid-career	Late career
Coefficient	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
All						
Fraction eligible	-0.0371*	-0.0029	0.0041	-0.0403**	-0.0003	0.0040
	(0.0207)	(0.0214)	(0.0202)	(0.0200)	(0.0203)	(0.0204)
Fraction eligible	0.0063	0.0365	0.0093	0.0031	0.0396	0.0090
+ Fraction eligible × Metropolitan	(0.0275)	(0.0253)	(0.0200)	(0.0264)	(0.0247)	(0.0195)
Low poverty						
Fraction eligible	-0.0440	-0.0093	-0.0186	-0.0512*	-0.0061	-0.0229
	(0.0285)	(0.0354)	(0.0279)	(0.0277)	(0.0337)	(0.0301)
Fraction eligible	-0.0280	0.0362	-0.0002	-0.0321	0.0389	-0.0048
+ Fraction eligible × Metropolitan	(0.0302)	(0.0319)	(0.0228)	(0.0284)	(0.0301)	(0.0245)
Moderate poverty						
Fraction eligible	-0.0171	0.0036	0.0186	-0.0156	-0.0028	0.0271
	(0.0321)	(0.0337)	(0.0303)	(0.0303)	(0.0310)	(0.0291)
Fraction eligible	0.0606*	0.0305	0.0351	0.0640**	0.0255	0.0424*
+ Fraction eligible × Metropolitan	(0.0301)	(0.0284)	(0.0237)	(0.0282)	(0.0264)	(0.0230)
High poverty						
Fraction eligible	-0.0507	-0.0402	0.0129	-0.0559	-0.0260	0.0131
	(0.0420)	(0.0495)	(0.0397)	(0.0405)	(0.0453)	(0.0368)
Fraction eligible	0.0037	0.0318	-0.0024	-0.0068	0.0495	-0.0039
+ Fraction eligible × Metropolitan	(0.0720)	(0.0524)	(0.0466)	(0.0683)	(0.0518)	(0.0433)
County fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
County controls	Yes	Yes	Yes	Yes	Yes	Yes
Observations	23,592	23,592	23,592	23,592	23,592	23,592

Notes: This table reports OLS and IV estimates of Equation (2) separately for early-career (at most 5 years in practice), mid-career (6-20 years in practice), and late-career (over 20 years in practice) OB/GYNs by tertiles of county poverty rates. The average poverty rates in 1980 and 1990 are used. The coefficients of interest are β_1 on the fraction eligible, and $\beta_1 + \beta_2$ on the sum of the fraction eligible and its interaction with the metropolitan indicator. Other regressors include county-specific characteristics (i.e., income per capita, unemployment rate, population density, fraction of 15-44-year-old female population, and fraction of black population), fee ratios, county fixed effects, and year fixed effects. Each column in the first panel corresponds to a separate regression. For each column in the remaining panels, a pooled model is estimated where every regressor is interacted with the poverty indicator. Standard errors in parentheses are clustered at the state level. Estimates are weighted by the county's female population, ages 15–44.
* $p < 0.10$ ** $p < 0.05$ *** $p < 0.01$

Table 4: Effect of eligibility changes on early-career OB/GYN supply by poverty and existing supply

Outcome: early-career OB/GYNs per 100,000				
	OLS estimation		IV estimation	
	Low supply	High supply	Low supply	High supply
Coefficient	**(1)**	**(2)**	**(3)**	**(4)**
Low poverty				
Fraction eligible	0.0376	-0.0562*	0.0296	-0.0643**
	(0.0358)	(0.0325)	(0.0315)	(0.0319)
Fraction eligible	-0.0149	-0.0291	-0.0222	-0.0332
+ Fraction eligible × Metropolitan	(0.0345)	(0.0318)	(0.0335)	(0.0298)
Moderate poverty				
Fraction eligible	0.0189	-0.0220	0.0119	-0.0207
	(0.0374)	(0.0354)	(0.0347)	(0.0335)
Fraction eligible	-0.0202	0.0730**	-0.0214	0.0765**
+ Fraction eligible × Metropolitan	(0.0528)	(0.0334)	(0.0494)	(0.0313)
High poverty				
Fraction eligible	0.0160	-0.0790	0.0104	-0.0844
	(0.0266)	(0.0567)	(0.0227)	(0.0539)
Fraction eligible	0.0468	-0.0275	0.0382	-0.0395
+ Fraction eligible × Metropolitan	(0.0453)	(0.0937)	(0.0392)	(0.0887)
County fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
County controls	Yes	Yes	Yes	Yes
Observations	11,997	11,595	11,997	11,595

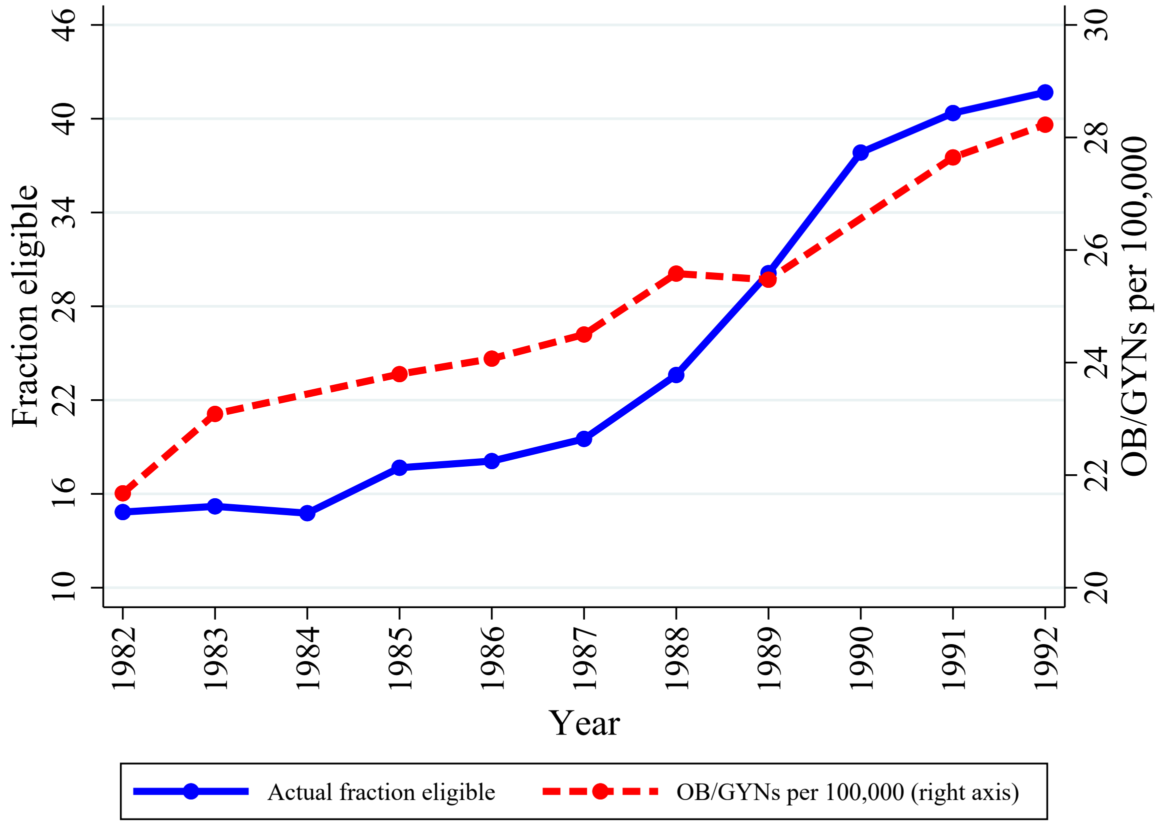
Notes: This table reports OLS and IV estimates of Equation (2) for early-career OB/GYNs by tertiles of county poverty rates. The average poverty rates in 1980 and 1990 are used. Estimates are reported separately for counties with high (i.e., above-median) versus low (i.e., below-median) non-early-career OB/GYN supply in the beginning of our sample period. The coefficients of interest are β_1 on the fraction eligible, and $\beta_1 + \beta_2$ on the sum of the fraction eligible and its interaction with the metropolitan indicator. Other regressors include county-specific characteristics (i.e., income per capita, unemployment rate, population density, fraction of 15-44-year-old female population, and fraction of black population), fee ratios, county fixed effects, and year fixed effects. For each column, a pooled model is estimated where every regressor is interacted with the poverty indicator. Standard errors in parentheses are clustered at the state level. Estimates are weighted by the county's female population, ages 15–44.
* $p < 0.10$ ** $p < 0.05$ *** $p < 0.01$

Table 5: Effect of eligibility changes on early-career OB/GYN supply by poverty and fee ratios

	Outcome: early-career OB/GYNs per 100,000			
	OLS estimation		IV estimation	
	Low fees	High fees	Low fees	High fees
Coefficient	**(1)**	**(2)**	**(3)**	**(4)**
Low poverty				
Fraction eligible	-0.0619	-0.0091	-0.0684*	-0.0198
	(0.0387)	(0.0414)	(0.0393)	(0.0387)
Fraction eligible	-0.0598	0.0191	-0.0624*	0.0105
+ Fraction eligible × Metropolitan	(0.0353)	(0.0475)	(0.0342)	(0.0441)
Moderate poverty				
Fraction eligible	-0.0428	0.0074	-0.0192	-0.0043
	(0.0441)	(0.0441)	(0.0423)	(0.0404)
Fraction eligible	0.0208	0.0975**	0.0443	0.0889**
+ Fraction eligible × Metropolitan	(0.0418)	(0.0388)	(0.0408)	(0.0366)
High poverty				
Fraction eligible	-0.0753	-0.0935	-0.0899	-0.0982
	(0.0891)	(0.0769)	(0.0841)	(0.0683)
Fraction eligible	0.0594	-0.0857	0.0398	-0.0980
+ Fraction eligible × Metropolitan	(0.0658)	(0.1237)	(0.0638)	(0.1152)
County fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
County controls	Yes	Yes	Yes	Yes
Observations	10,388	13,204	10,388	13,204

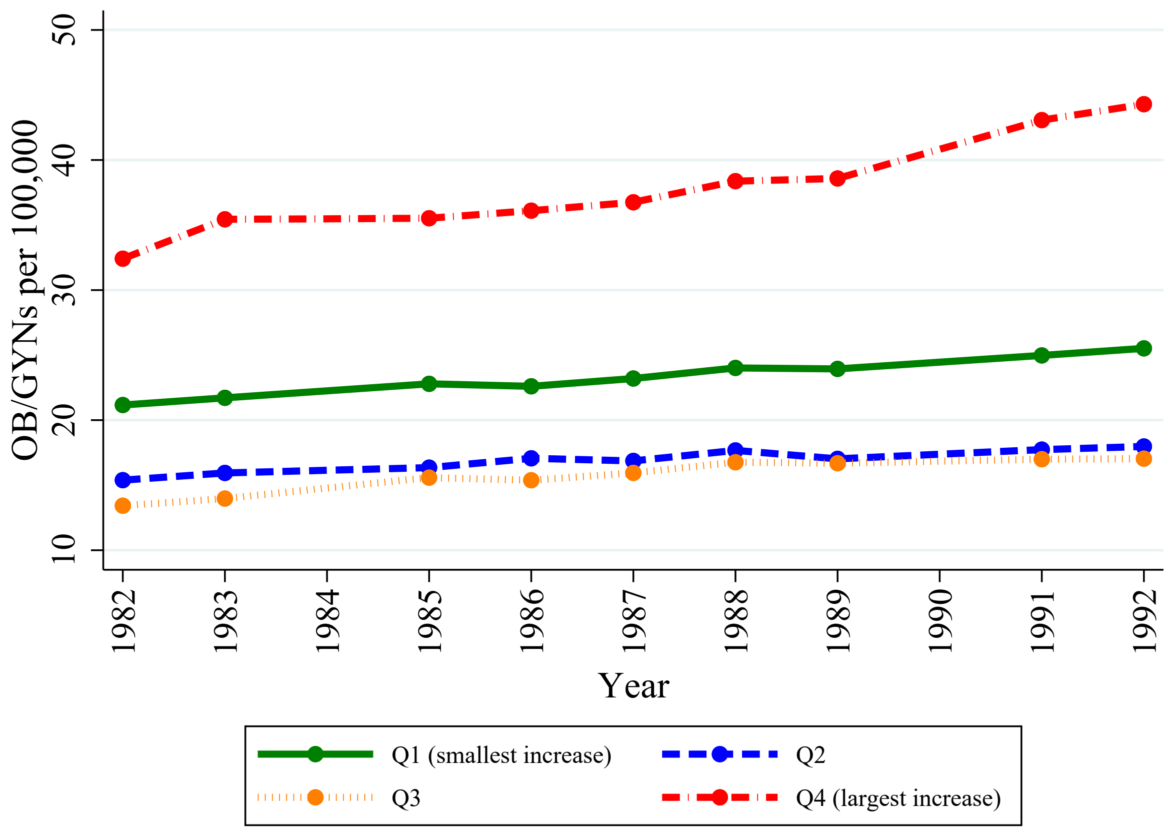
Notes: This table reports OLS and IV estimates of Equation (2) for early-career OB/GYNs by tertiles of county poverty rates. The average poverty rates in 1980 and 1990 are used. Estimates are reported separately for counties in states with high (i.e., above-median) versus low (i.e., below-median) Medicaid-to-private fee ratios in the beginning of our sample period. The coefficients of interest are β_1 on the fraction eligible, and $\beta_1 + \beta_2$ on the sum of the fraction eligible and its interaction with the metropolitan indicator. Other regressors include county-specific characteristics (i.e., income per capita, unemployment rate, population density, fraction of 15-44-year-old female population, and fraction of black population), fee ratios, county fixed effects, and year fixed effects. For each column, a pooled model is estimated where every regressor is interacted with the poverty indicator. Standard errors in parentheses are clustered at the state level. Estimates are weighted by the county's female population, ages 15–44.
* $p < 0.10$ ** $p < 0.05$ *** $p < 0.01$

Figure 1: Trends in fraction eligible and OB/GYN supply



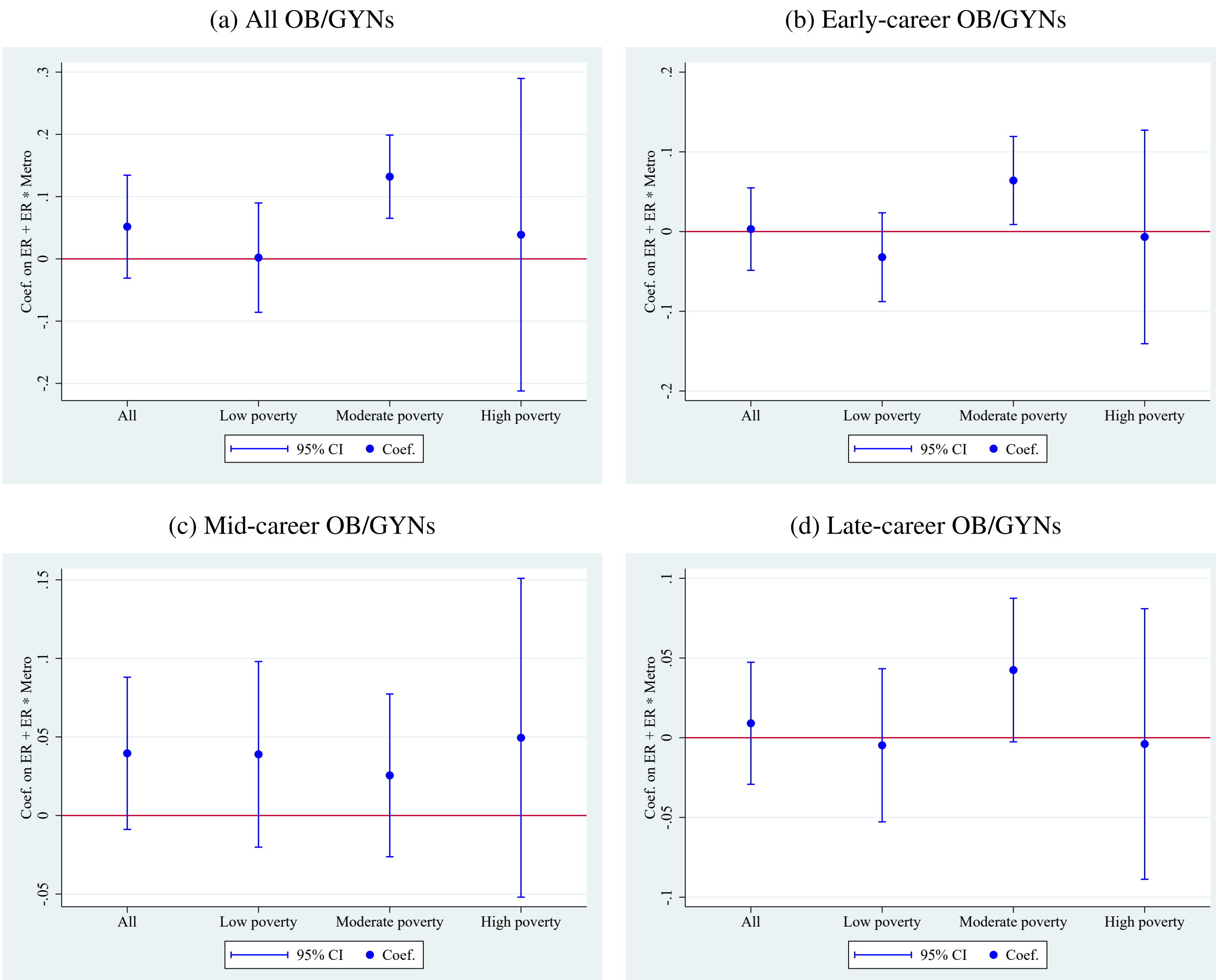
Notes: This figure shows aggregate trends in the actual fraction eligible (left y-axis) and OB/GYNs per 100,000 females aged 14–55 (right y-axis) during 1982–1992. Data on OB/GYN counts are missing in 1984 and 1990.

Figure 2: Trends in OB/GYN supply, by size of of the eligibility expansions



Notes: This figure shows aggregate trends in OB/GYNs per 100,000 females aged 14–55, by the percent increase in eligibility rates between 1982 and 1992. Data on OB/GYN counts are missing in 1984 and 1990.

Figure 3: Coefficient plots of the effect of eligibility changes on OB/GYN supply by poverty for metropolitan counties



Notes: This figure plots IV estimates of the sum of the fraction eligible and its interaction with the metropolitan indicator from Equation (2) by tertiles of county poverty rates for all OB/GYNs in Panel (a) and for OB/GYNs by career stage in Panels (b)-(d). The average poverty rates in 1980 and 1990 are used. Bands indicate 95% confidence intervals. Standard errors are clustered at the state level. Estimates are weighted by the county's female population, ages 15–44.

Appendix

Table A1: Fraction eligible and OB/GYN supply

Year	Actual fraction eligible	OB/GYNs per 100,000
1982	14.84	21.68
1983	15.20	23.09
1984	14.77	
1985	17.67	23.79
1986	18.10	24.07
1987	19.52	24.50
1988	23.61	25.58
1989	30.12	25.47
1990	37.84	
1991	40.36	27.65
1992	41.68	28.23

Notes: This table shows the actual fraction eligible and OB/GYNs per 100,000 females aged 14–55 by year for 1982–1992. Data on OB/GYN counts are missing in 1984 and 1990.

Table A2: First-stage results for correlation between actual eligibility rates and simulated eligibility rates

Outcome: Actual fraction eligible

	First-stage estimation			
Coefficient	**(1)**	**(2)**	**(3)**	**(4)**
Simulated fraction eligible	0.8717***	0.8734***	0.8726***	0.8761***
	(0.0206)	(0.0219)	(0.0188)	(0.0194)
Simulated fraction eligible + Simulated fraction eligible × Metropolitan		0.8703*** (0.0200)		0.8705*** (0.0187)
County fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
County controls	No	No	Yes	Yes
Observations	23,592	23,592	23,592	23,592
Partial *F* -statistic	1,794	944	2,156	1,099

Notes: This table reports first-stage estimates of Equations (1) and (2) using the actual fraction eligible as the outcome. The coefficients of interest are β_1 on the simulated fraction eligible, and $\beta_1 + \beta_2$ on the sum of the simulated fraction eligible and its interaction with the metropolitan indicator (for Equation 2). Columns (1) and (2) include only county fixed effects and year fixed effects. Columns (3) and (4) also include county-specific characteristics (i.e., income per capita, unemployment rate, population density, fraction of 15-44-year-old female population, and fraction of black population) and fee ratios. Each column corresponds to a separate regression. Partial F-statistic is reported for each regression. Standard errors in parentheses are clustered at the state level. Estimates are weighted by the county's female population, ages 15–44.

* $p < 0.10$ ** $p < 0.05$ *** $p < 0.01$

Table A3: Effect of eligibility changes on early-career OB/GYN supply by poverty

Outcome: early-career OB/GYNs per 100,000				
	OLS estimation		IV estimation	
Coefficient	**(1)**	**(2)**	**(3)**	**(4)**
All				
Fraction eligible	-0.0106	-0.0371*	-0.0136	-0.0403**
	(0.0214)	(0.0207)	(0.0209)	(0.0200)
Fraction eligible		0.0063		0.0031
+ Fraction eligible × Metropolitan		(0.0275)		(0.0264)
Low poverty				
Fraction eligible	-0.0303	-0.0440	-0.0351	-0.0512*
	(0.0282)	(0.0285)	(0.0267)	(0.0277)
Fraction eligible		-0.0280		-0.0321
+ Fraction eligible × Metropolitan		(0.0302)		(0.0284)
Moderate poverty				
Fraction eligible	0.0446	-0.0171	0.0463*	-0.0156
	(0.0289)	(0.0321)	(0.0271)	(0.0303)
Fraction eligible		0.0606*		0.0640**
+ Fraction eligible × Metropolitan		(0.0301)		(0.0282)
High poverty				
Fraction eligible	-0.0340	-0.0507	-0.0401	-0.0559
	(0.0463)	(0.0420)	(0.0445)	(0.0405)
Fraction eligible		0.0037		-0.0068
+ Fraction eligible × Metropolitan		(0.0720)		(0.0683)
County fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
County controls	Yes	Yes	Yes	Yes
Observations	23,592	23,592	23,592	23,592

Notes: This table reports OLS and IV estimates of Equations (1) and (2) for early-career OB/GYNs by tertiles of county poverty rates. The average poverty rates in 1980 and 1990 are used. The coefficients of interest are β_1 on the fraction eligible, and $\beta_1 + \beta_2$ on the sum of the fraction eligible and its interaction with the metropolitan indicator (for Equation 2). Other regressors include county-specific characteristics (i.e., income per capita, unemployment rate, population density, fraction of 15-44-year-old female population, and fraction of black population), fee ratios, county fixed effects, and year fixed effects. Each column in the first panel corresponds to a separate regression. For each column in the remaining panels, a pooled model is estimated where every regressor is interacted with the poverty indicator. Standard errors in parentheses are clustered at the state level. Estimates are weighted by the county's female population, ages 15–44.

* $p < 0.10$ ** $p < 0.05$ *** $p < 0.01$

Table A4: Effect of eligibility changes on mid-career OB/GYN supply by poverty

Outcome: mid-career OB/GYNs per 100,000				
	OLS estimation		IV estimation	
Coefficient	**(1)**	**(2)**	**(3)**	**(4)**
All				
Fraction eligible	0.0212	-0.0029	0.0242	-0.0003
	(0.0195)	(0.0214)	(0.0191)	(0.0203)
Fraction eligible		0.0365		0.0396
+ Fraction eligible × Metropolitan		(0.0253)		(0.0247)
Low poverty				
Fraction eligible	0.0296	-0.0093	0.0319	-0.0061
	(0.0310)	(0.0354)	(0.0292)	(0.0337)
Fraction eligible		0.0362		0.0389
+ Fraction eligible × Metropolitan		(0.0319)		(0.0301)
Moderate poverty				
Fraction eligible	0.0249	0.0036	0.0192	-0.0028
	(0.0266)	(0.0337)	(0.0244)	(0.0310)
Fraction eligible		0.0305		0.0255
+ Fraction eligible × Metropolitan		(0.0284)		(0.0264)
High poverty				
Fraction eligible	-0.0181	-0.0402	-0.0017	-0.0260
	(0.0417)	(0.0495)	(0.0395)	(0.0453)
Fraction eligible		0.0318		0.0495
+ Fraction eligible × Metropolitan		(0.0524)		(0.0518)
County fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
County controls	Yes	Yes	Yes	Yes
Observations	23,592	23,592	23,592	23,592

Notes: This table reports OLS and IV estimates of Equations (1) and (2) for mid-career OB/GYNs by tertiles of county poverty rates. The average poverty rates in 1980 and 1990 are used. The coefficients of interest are β_1 on the fraction eligible, and $\beta_1 + \beta_2$ on the sum of the fraction eligible and its interaction with the metropolitan indicator (for Equation 2). Other regressors include county-specific characteristics (i.e., income per capita, unemployment rate, population density, fraction of 15-44-year-old female population, and fraction of black population), fee ratios, county fixed effects, and year fixed effects. Each column in the first panel corresponds to a separate regression. For each column in the remaining panels, a pooled model is estimated where every regressor is interacted with the poverty indicator. Standard errors in parentheses are clustered at the state level. Estimates are weighted by the county's female population, ages 15–44.

* $p < 0.10$ ** $p < 0.05$ *** $p < 0.01$

Table A5: Effect of eligibility changes on late-career OB/GYN supply by poverty

Outcome: late-career OB/GYNs per 100,000				
	OLS estimation		IV estimation	
Coefficient	**(1)**	**(2)**	**(3)**	**(4)**
All				
Fraction eligible	0.0073	0.0041	0.0071	0.0040
	(0.0190)	(0.0202)	(0.0190)	(0.0204)
Fraction eligible		0.0093		0.0090
+ Fraction eligible × Metropolitan		(0.0200)		(0.0195)
Low poverty				
Fraction eligible	-0.0029	-0.0186	-0.0076	-0.0229
	(0.0221)	(0.0279)	(0.0242)	(0.0301)
Fraction eligible		-0.0002		-0.0048
+ Fraction eligible × Metropolitan		(0.0228)		(0.0245)
Moderate poverty				
Fraction eligible	0.0317	0.0186	0.0390*	0.0271
	(0.0224)	(0.0303)	(0.0220)	(0.0291)
Fraction eligible		0.0351		0.0424*
+ Fraction eligible × Metropolitan		(0.0237)		(0.0230)
High poverty				
Fraction eligible	0.0082	0.0129	0.0077	0.0131
	(0.0411)	(0.0397)	(0.0382)	(0.0368)
Fraction eligible		-0.0024		-0.0039
+ Fraction eligible × Metropolitan		(0.0466)		(0.0433)
County fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
County controls	Yes	Yes	Yes	Yes
Observations	23,592	23,592	23,592	23,592

Notes: This table reports OLS and IV estimates of Equations (1) and (2) for late-career OB/GYNs by tertiles of county poverty rates. The average poverty rates in 1980 and 1990 are used. The coefficients of interest are β_1 on the fraction eligible, and $\beta_1 + \beta_2$ on the sum of the fraction eligible and its interaction with the metropolitan indicator (for Equation 2). Other regressors include county-specific characteristics (i.e., income per capita, unemployment rate, population density, fraction of 15-44-year-old female population, and fraction of black population), fee ratios, county fixed effects, and year fixed effects. Each column in the first panel corresponds to a separate regression. For each column in the remaining panels, a pooled model is estimated where every regressor is interacted with the poverty indicator. Standard errors in parentheses are clustered at the state level. Estimates are weighted by the county's female population, ages 15–44.

* $p < 0.10$ ** $p < 0.05$ *** $p < 0.01$

Table A6: Effect of eligibility changes on early-career OB/GYN supply by poverty and adoption of tort reforms

Outcome: early-career OB/GYNs per 100,000				
	OLS estimation		IV estimation	
	No tort reforms	Tort reforms	No tort reforms	Tort reforms
Coefficient	**(1)**	**(2)**	**(3)**	**(4)**
Low poverty				
Fraction eligible	-0.0567	-0.0407	-0.0613*	-0.0509
	(0.0365)	(0.0350)	(0.0361)	(0.0330)
Fraction eligible	-0.0077	-0.0392	-0.0082	-0.0482*
+ Fraction eligible × Metropolitan	(0.0455)	(0.0297)	(0.0424)	(0.0271)
Moderate poverty				
Fraction eligible	-0.0796	0.0125	-0.0838*	0.0146
	(0.0514)	(0.0279)	(0.0479)	(0.0254)
Fraction eligible	0.0329	0.0683**	0.0329	0.0707***
+ Fraction eligible × Metropolitan	(0.0517)	(0.0277)	(0.0474)	(0.0266)
High poverty				
Fraction eligible	-0.0393	-0.0519	-0.0413	-0.0679
	(0.0532)	(0.0571)	(0.0518)	(0.0576)
Fraction eligible	-0.0037	-0.0206	-0.0160	-0.0352
+ Fraction eligible × Metropolitan	(0.1190)	(0.0531)	(0.1162)	(0.0523)
County fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
County controls	Yes	Yes	Yes	Yes
Observations	10,388	13,204	10,388	13,204

Notes: This table reports OLS and IV estimates of Equation (2) for early-career OB/GYNs by tertiles of county poverty rates. The average poverty rates in 1980 and 1990 are used. Estimates are reported separately for counties in states with and without tort reform in place, following Malani and Reif (2015): collateral source reform, joint and several liability reform, non-economic damage cap, punitive damage cap, or split recovery reform. The coefficients of interest are β_1 on the fraction eligible, and $\beta_1 + \beta_2$ on the sum of the fraction eligible and its interaction with the metropolitan indicator. Other regressors include county-specific characteristics (i.e., income per capita, unemployment rate, population density, fraction of 15-44-year-old female population, and fraction of black population), fee ratios, county fixed effects, and year fixed effects. For each column, a pooled model is estimated where every regressor is interacted with the poverty indicator. Standard errors in parentheses are clustered at the state level. Estimates are weighted by the county's female population, ages 15–44.

* $p < 0.10$ ** $p < 0.05$ *** $p < 0.01$

Table A7: Effect of eligibility changes on early-career OB/GYN supply by poverty, excluding county controls and fee ratios

Outcome: early-career OB/GYNs per 100,000		
	OLS estimation	IV estimation
Coefficient	(1)	(2)
Low poverty		
Fraction eligible	-0.0418	-0.0487
	(0.0334)	(0.0326)
Fraction eligible	-0.0336	-0.0376
+ Fraction eligible × Metropolitan	(0.0336)	(0.0315)
Moderate poverty		
Fraction eligible	-0.0153	-0.0142
	(0.0392)	(0.0375)
Fraction eligible	0.0582*	0.0616**
+ Fraction eligible × Metropolitan	(0.0329)	(0.0310)
High poverty		
Fraction eligible	-0.0292	-0.0337
	(0.0322)	(0.0303)
Fraction eligible	0.0202	0.0114
+ Fraction eligible × Metropolitan	(0.0616)	(0.0577)
County fixed effects	Yes	Yes
Year fixed effects	Yes	Yes
County controls	No	No
Observations	23,592	23,592

Notes: This table reports OLS and IV estimates of Equation (2) for early-career OB/GYNs by tertiles of county poverty rates. The average poverty rates in 1980 and 1990 are used. The coefficients of interest are β_1 on the fraction eligible, and $\beta_1 + \beta_2$ on the sum of the fraction eligible and its interaction with the metropolitan indicator. County fixed effects and year fixed effects are included. For each column, a pooled model is estimated where every regressor is interacted with the poverty indicator. Standard errors in parentheses are clustered at the state level. Estimates are weighted by the county's female population, ages 15–44.

* $p < 0.10$ ** $p < 0.05$ *** $p < 0.01$

Table A8: Effect of eligibility changes on early-career OB/GYN supply by poverty, unweighted

Outcome: early-career OB/GYNs per 100,000		
	OLS estimation	**IV estimation**
Coefficient	**(1)**	**(2)**
Low poverty		
Fraction eligible	0.0369	0.0293
	(0.0263)	(0.0249)
Fraction eligible	0.0241	0.0181
+ Fraction eligible × Metropolitan	(0.0248)	(0.0241)
Moderate poverty		
Fraction eligible	0.0063	0.0018
	(0.0300)	(0.0281)
Fraction eligible	0.0562*	0.0539*
+ Fraction eligible × Metropolitan	(0.0305)	(0.0297)
High poverty		
Fraction eligible	-0.0910	-0.0911
	(0.0821)	(0.0732)
Fraction eligible	-0.0870	-0.0907
+ Fraction eligible × Metropolitan	(0.1337)	(0.1235)
County fixed effects	Yes	Yes
Year fixed effects	Yes	Yes
County controls	Yes	Yes
Observations	23,592	23,592

Notes: This table reports OLS and IV estimates of Equation (2) for early-career OB/GYNs by tertiles of county poverty rates. The average poverty rates in 1980 and 1990 are used. The coefficients of interest are β_1 on the fraction eligible, and $\beta_1 + \beta_2$ on the sum of the fraction eligible and its interaction with the metropolitan indicator. Other regressors include county-specific characteristics (i.e., income per capita, unemployment rate, population density, fraction of 15-44-year-old female population, and fraction of black population), fee ratios, county fixed effects, and year fixed effects. For each column, a pooled model is estimated where every regressor is interacted with the poverty indicator. Standard errors in parentheses are clustered at the state level. Estimates are unweighted.
* $p < 0.10$ ** $p < 0.05$ *** $p < 0.01$

Table A9: Effect of eligibility changes on early-career OB/GYN supply by poverty, using non-lagged eligibility rates

Outcome: early-career OB/GYNs per 100,000		
	OLS estimation	**IV estimation**
Coefficient	**(1)**	**(2)**
Low poverty		
Fraction eligible	-0.0272	-0.0340
	(0.0268)	(0.0267)
Fraction eligible	-0.0177	-0.0236
+ Fraction eligible × Metropolitan	(0.0285)	(0.0279)
Moderate poverty		
Fraction eligible	-0.0344	-0.0207
	(0.0305)	(0.0281)
Fraction eligible	0.0424	0.0563**
+ Fraction eligible × Metropolitan	(0.0285)	(0.0266)
High poverty		
Fraction eligible	-0.0365	-0.0363
	(0.0389)	(0.0368)
Fraction eligible	-0.0055	-0.0055
+ Fraction eligible × Metropolitan	(0.0661)	(0.0623)
County fixed effects	Yes	Yes
Year fixed effects	Yes	Yes
County controls	Yes	Yes
Observations	23,592	23,592

Notes: This table reports OLS and IV estimates of Equation (2) for early-career OB/GYNs by tertiles of county poverty rates. The average poverty rates in 1980 and 1990 are used. The actual, non-lagged eligibility rates are used. The coefficients of interest are β_1 on the fraction eligible, and $\beta_1 + \beta_2$ on the sum of the fraction eligible and its interaction with the metropolitan indicator. Other regressors include county-specific characteristics (i.e., income per capita, unemployment rate, population density, fraction of 15-44-year-old female population, and fraction of black population), fee ratios, county fixed effects, and year fixed effects. For each column, a pooled model is estimated where every regressor is interacted with the poverty indicator. Standard errors in parentheses are clustered at the state level. Estimates are weighted by the county's female population, ages 15–44.

* $p < 0.10$ ** $p < 0.05$ *** $p < 0.01$

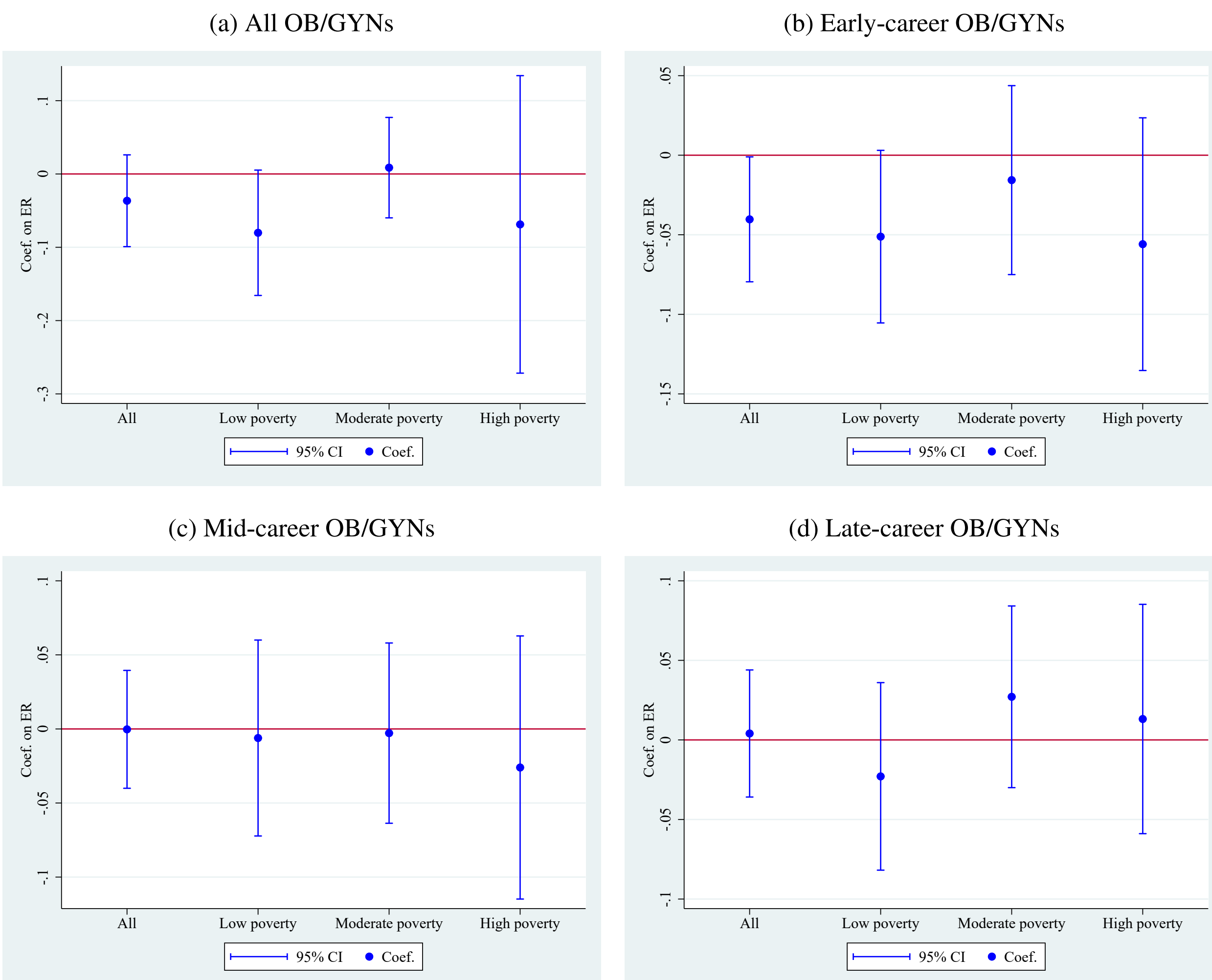
Table A10: Effect of eligibility changes on early-career OB/GYN supply by alternative poverty groups

Outcome: early-career OB/GYNs per 100,000				
	OLS estimation		IV estimation	
Coefficient	**(1)**	**(2)**	**(3)**	**(4)**
Quartile 1 (lowest poverty)				
Fraction eligible	-0.0459	-0.0514	-0.0513*	-0.0601
	(0.0321)	(0.0401)	(0.0311)	(0.0389)
Fraction eligible	-0.0299	-0.0401	-0.0333	-0.0458
+ Fraction eligible × Metropolitan	(0.0295)	(0.0306)	(0.0279)	(0.0299)
Quartile 2				
Fraction eligible	-0.0490	-0.0662*	-0.0472	-0.0620*
	(0.0304)	(0.0389)	(0.0289)	(0.0375)
Fraction eligible	-0.0101	-0.0022	-0.0051	0.0056
+ Fraction eligible × Metropolitan	(0.0344)	(0.0478)	(0.0316)	(0.0442)
Quartile 3				
Fraction eligible	0.0087	0.0257	0.0007	0.0264
	(0.0397)	(0.0351)	(0.0382)	(0.0329)
Fraction eligible	0.0935**	0.0617*	0.0858**	0.0637*
+ Fraction eligible × Metropolitan	(0.0392)	(0.0355)	(0.0374)	(0.0342)
Quintile 4				
Fraction eligible	-0.0590	-0.0227	-0.0253	-0.0195
	(0.0563)	(0.0457)	(0.0343)	(0.0297)
Fraction eligible	0.0260	0.0709	0.0711	0.0785**
+ Fraction eligible × Metropolitan	(0.0818)	(0.0517)	(0.0526)	(0.0394)
Quintile 5 (highest poverty)				
Fraction eligible		-0.1271*		-0.0594
		(0.0712)		(0.0392)
Fraction eligible + Fraction		-0.1060		-0.0122
eligible X Metropolitan		(0.1002)		(0.0418)
County fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
County controls	Yes	Yes	Yes	Yes
Observations	23,592	23,592	23,592	23,592

Notes: This table reports OLS and IV estimates of Equation (2) for early-career OB/GYNs by quartiles and quintiles of county poverty rates. The average poverty rates in 1980 and 1990 are used. The coefficients of interest are β_1 on the fraction eligible, and $\beta_1 + \beta_2$ on the sum of the fraction eligible and its interaction with the metropolitan indicator. Other regressors include county-specific characteristics (i.e., income per capita, unemployment rate, population density, fraction of 15-44-year-old female population, and fraction of black population), fee ratios, county fixed effects, and year fixed effects. For each column, a pooled model is estimated where every regressor is interacted with the poverty indicator. Standard errors in parentheses are clustered at the state level. Estimates are weighted by the county's female population, ages 15–44.

* $p < 0.10$ ** $p < 0.05$ *** $p < 0.01$

Figure A1: Coefficient plots of the effect of eligibility changes on OB/GYN supply by poverty for non-metropolitan counties



Notes: This figure plots IV estimates of the fraction eligible from Equation (2) by tertiles of county poverty rates for all OB/GYNs in Panel (a) and for OB/GYNs by career stage in Panels (b)-(d). The average poverty rates in 1980 and 1990 are used. Bands indicate 95% confidence intervals. Standard errors are clustered at the state level. Estimates are weighted by the county's female population, ages 15–44.