

# The Impact of a \$10,000 Bonus on Special Education Teacher Shortages in Hawai'i

Roddy Theobald

Zeyu Xu

Allison Gilmour

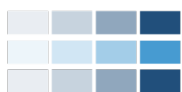
Lisa Lachlan-Hache

Liz Bettini

Nathan Jones

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**Roddy Theobald**

*American Institutes for Research / CALDER*

**Zeyu Xu**

*American Institutes for Research / CALDER*

**Allison Gilmour**

*American Institutes for Research*

**Lisa Lachlan-Hache**

*American Institutes for Research / GTL Center*

**Liz Bettini**

*Boston University*

**Nathan Jones**

*Boston University*

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CALDER • American Institutes for Research  
1400 Crystal Drive 10<sup>th</sup> Floor, Arlington, VA 22202  
202-403-5796 • [www.caldercenter.org](http://www.caldercenter.org)

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Roddy Theobald, Zeyu Xu, Allison Gilmour, Lisa Lachlan-Hache, Liz Bettini, & Nathan Jones

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**Abstract**

We study the impact of a bonus policy implemented by Hawai‘i Public Schools starting in fall 2020 that raised the salaries of all special education teachers in the state by \$10,000. We estimate that the introduction of this policy reduced the proportion of vacant special education teaching positions by 32%, or 1.2 percentage points, and the proportion of special education positions that were vacant or filled by an unlicensed teacher by 35%, or 4.0 percentage points. The bonus policy did not have significant impacts on special education teacher retention; instead, the impacts of the policy were driven almost entirely by an increase in the number of general education teachers in the state who moved into open special education teaching positions. The effects of the bonus policy were also largest in historically hard-to-staff schools in which all teachers also received “tiered school” bonuses of up to \$8,000. Hawai‘i therefore represents a unique but instructive case of how strategic financial incentives can help address special education teacher shortages.

## 1. Introduction

There is a crisis in special education: in many parts of the country, there are simply not enough qualified special educators to meet the needs of students with disabilities in K–12 public schools. This issue is not new; school systems have struggled to staff special education for decades (Billingsley, 1993; Carriker, 1989; Cowan et al., 2016; Mason-Williams et al., 2020). What is new is that more states and districts are offering strategic financial incentives to special educators with the goal of proactively addressing special education teacher shortages (Putnam & Gerber, 2022).

One of the highest-profile recent examples of this trend is in Hawai‘i Public Schools, which implemented a bonus policy starting in fall 2020 that raised the salaries of all special education teachers in the state by \$10,000 (McCoy, 2022). This special education bonus, combined with additional bonuses for teachers working in historically hard-to-staff schools, means that special education teachers in the hardest-to-staff schools could be making up to \$18,000 more per year than prior to the bonus policy. Although prior evidence on teacher financial incentives (discussed in the next section) suggests that this policy *could* move the needle in terms of addressing long-standing special education teacher shortages in the state, there is no existing causal evidence on the impact of bonuses specifically offered to special education teachers on these critical staffing challenges.

We therefore provide the first large-scale plausibly causal evidence on this topic by studying the impact of this bonus policy on special education teacher shortages using longitudinal position-level staffing data from Hawai‘i Public Schools from the 2014–15 through the 2022–23 school years. The Hawai‘i data are unique in that they provide an annual snapshot from October 1 of each school year of every school staff position in the state *whether or not the position was filled*, along with indicators for whether positions that are filled are filled by a

teacher who meets the state's licensure requirements. This allows us to create two position-level measures of staffing challenges (positions that are entirely vacant and/or positions that are filled by unlicensed teachers) in addition to annual indicators for the attrition and hiring of individual teachers in the state. We use these data to investigate three research questions:

1. Research Question 1: What is the average causal effect of the \$10,000 bonus policy on special education teacher shortages in Hawai'i, as measured by the proportion of positions that are vacant and/or by the proportion of positions filled by unlicensed teachers?
2. Research Question 2: To what extent are these effects driven by changes in patterns of teacher attrition and movement between special education and general education teaching positions?
3. Research Question 3: How does the effect of the \$10,000 bonus policy vary between hard-to-staff schools in which all teachers received additional bonuses and other schools?

We investigate Research Question (RQ) 1 by estimating difference-in-differences (DID) models that exploit the fact that an untreated group of teachers (general education teachers) were ineligible for the \$10,000 special education teacher bonus. Our DID estimates suggest that the introduction of the special education teacher bonus policy reduced the proportion of vacant special education teaching positions by 32%, or 1.2 percentage points, and the proportion of special education positions that were vacant or filled by an unlicensed teacher by 35%, or 4.0 percentage points. These findings are robust to a variety of model specifications that make comparisons within schools, within schools and position types, and within schools and school years. We also fail to reject the assumption of parallel trends between treated and control

teachers, even when using an estimator that is robust to heterogeneous treatment effects across groups and over time (Callaway & Sant’Anna, 2021). We therefore interpret these estimates as the average causal effects of the special education bonus policy on these proxies for special education teacher shortages, and we conclude that the \$10,000 bonus was sufficiently large to significantly move the needle in terms of special education teacher shortages in the state.

When we explore potential mechanisms for these effects as part of RQ2, we find, perhaps surprisingly, that the bonus policy had no significant impact on special education teacher retention. Instead, we document a large increase in the proportion of open special education positions that were filled by general education teachers in the state after the introduction of the policy. Although descriptive, this suggests that the policy impacted special education shortages in the short term through a redistribution of existing teachers, rather than by retaining more special education teachers within the school system. Future work will be necessary to determine the long-term impacts of the policy, including whether the bonuses induced more potential teachers to pursue a special education teaching position in the first place.

Finally, we estimate triple difference (DDD) models to explore heterogeneity in the impacts of the bonus policy across different “tiers” of schools that were identified for additional schoolwide bonuses under the state’s bonus policy due to a variety of factors including prior vacancy rates. For example, all teachers in Tier 4 schools received bonuses of \$8,000 under the policy, which meant that special education teachers in Tier 4 schools received \$18,000 bonuses, whereas general education teachers in the same school received \$8,000 bonuses. Estimates from the DDD models suggest that the impact of the bonus policy on special education vacancies was driven almost entirely by impacts in these Tier 4 schools, whereas the impact of the bonus policy on positions that were vacant or filled by unlicensed teachers was significant across the different



school tiers but larger in Tier 3 and Tier 4 schools. This suggests that the bonus policy was differentially effective at addressing special education teacher shortages in historically hard-to-staff schools.

## **2. Literature Review**

We believe this to be the first causal study of financial incentives specifically targeted to special education teachers, even as states and districts have increasingly looked to use bonuses or base pay raises to address special education teacher shortages (Gilmour et al., 2023; Putnam & Gerber, 2022). There is a well-established literature base showing that bonuses can help fill empty teaching positions and can improve teacher retention, but this work disproportionately comes from studies about general education teachers or programs targeting other hard-to-staff subjects such as math and science. We begin this section by reviewing this existing literature, and then we describe the specific elements that may result in special educators responding differently to bonuses than their counterparts in general education or other hard-to-staff areas.

### ***2.1 Empirical Evidence on Teacher Bonuses***

States and districts have implemented bonus programs to address staffing challenges which have targeted: teachers in hard-to-staff subjects (Clotfelter et al., 2008a, b; Feng & Sass, 2018), particularly effective teachers (Dee & Wyckoff, 2015; Springer et al., 2016), teachers in hard-to-staff schools (Castro & Esposito, 2021; Cowan & Goldhaber, 2018; Glazerman et al., 2013), or some combination of these teacher populations. Clotfelter et al. (2005, 2008a, 2008b) investigated the effects of a \$1,800 annual bonus in North Carolina to math, science, and special education teachers in middle and high schools, in which 80% or more of the students qualified for free or reduced-price lunch or 50% or more students scored below grade level in Algebra 1 or Biology. They did not find any immediate effects on teacher retention, due in large part to implementation challenges. Primarily, principals did not know teachers in their school were

eligible for the bonus, thus teachers did not know they were eligible for the bonus. However, in their follow-up study, Clotfelter et al. (2008a, 2008b) used discrete time hazard models to examine long-term retention due to the bonus program and found the bonus resulted in a 15%–18% reduction in the probability of teacher turnover. These reductions in the probability of turnover were entirely driven by math teachers. Feng and Sass (2018) examined the impact of a similar program in Florida that provided \$1,200 bonuses to middle or high school teachers in hard-to-staff areas (science, math, foreign language, and special education) using a difference-in-difference design, with a triple difference because of a co-occurring loan forgiveness program. The bonus was associated with a 32.2% decline in the probability that a teacher left teaching in Florida.

Bonuses have also been used to retain effective teachers (Dee & Wyckoff, 2015), particularly in hard-to-staff schools (Springer et al., 2016). In a study of a bonus program tied to teacher effectiveness in Washington, DC, Dee and Wyckoff (2015) found that teachers who received a highly effective rating accompanied by a bonus were three percentage points more likely to remain in their school than teachers who just missed the rating qualifying them for a bonus, but the difference was not statistically significant. Springer and colleagues (2016) investigated the effect of a \$5,000 retention bonus in Tennessee tied to teachers' evaluation scores and teaching in low-performing schools using a regression discontinuity design. Although there were no overall effects of the bonus, the bonus increased the probability of retention for teachers in tested subjects and grades by 20%. Similar to the bonus program in North Carolina, there were implementation challenges in this program including eligible teachers not receiving the bonus and ineligible teachers receiving the bonus.

Bonus programs can also incentivize the transfer of effective teachers into harder to staff schools (Glazerman et al., 2013) and the retention of teachers in disadvantaged schools (Castro & Esposito, 2021; Cowan & Goldhaber, 2018; Elacqua et al., 2022). In a randomized control trial in 10 districts across seven states, Glazerman and colleagues examined the effect of a novel transfer incentive bonus: Teachers in the top 20% of value-added in their subject in the state assigned to the treatment group were eligible for a \$20,000 bonus distributed across two years if they moved into a school serving higher proportions of disadvantaged students. The bonus resulted in improved retention for the teachers who took the bonus, with 93% of teachers who received the bonus staying in their positions compared to 70% of teachers in the control group. However, there was only an effect of the bonus on retention while the teachers received the bonus, suggesting the need for ongoing bonuses instead of short-term bonuses. Cowan and Goldhaber (2018) studied a bonus program in Washington for National Board certified teachers to work in schools serving higher proportions of students qualifying for free/reduced lunch. Using a regression discontinuity design, they estimated that school eligibility for the bonus program was associated with a 3.2 to 4.3 percentage point decline in the probability that targeted teachers left the school.

These findings are bolstered by two recent studies from outside the United States. In a study of a bonus program in Chile also using a regression discontinuity design, Elacqua et al. (2022) found that monthly bonuses, corresponding to 5%–16% of teachers' salaries with additional increases for teaching in schools with more than 60% of the students from low-income backgrounds, increased the probability that an effective teacher worked at a disadvantaged schools by 17 to 21 percentage points two years after the bonus was implemented. The bonus did not improve retention of effective teachers in more advantaged schools. In Peru, Castro and

Esposito (2021) similarly found monthly bonuses of about 26% of a beginning teacher's salary for teachers in rural schools was associated with a 1.6 to 3.4 percentage point increase in the probability of filling vacancies in the schools. The bonuses were associated with 1.5 to 4.9 percentage point reductions in attrition. An unintentional consequence of the increase in recruitment and retention in rural schools were more unfilled vacancies in nearby schools in which teachers did not qualify for the bonus. This set of studies provides strong support for the use of on-going bonuses to retain teachers in hard-to-staff settings.

## **2.2 Bonus Evidence Related to Special Education**

The emerging evidence discussed in Section 2.1 that focuses on teachers in tested grades and subjects (Springer et al., 2016), teachers in certain fields (Clotfelter et al., 2008a, 2008b; Feng & Sass, 2018), and teachers in harder-to-staff schools (Castro & Esposito, 2021; Cowan & Goldhaber, 2018; Elacqua et al., 2022; Glazerman et al., 2013) suggests that bonuses can improve teacher retention and decrease teacher vacancies. Although most of these studies have not focused specifically on special education teachers, the subset that focused more broadly on “hard-to-staff subjects” either did not find significant effects or did not disaggregate results for special education teachers. For example, when Clotfelter et al. (2008a, 2008b) disaggregated the effects of bonuses by teacher subject, they found no significant effect on special education teachers, whereas Feng and Sass (2018) did not disaggregate results for special educators.

The failure to find a statistically significant association between bonuses and special educator retention could reflect sample size limitations, but it could also reflect the specifics of special educators' roles in schools. Three important differences between special education and other teaching positions in public schools inform this study. First, although bonuses are often framed as a mechanism for replacing financial gains lost by becoming an educator instead of

another career using the same education and skills—and although this idea undergirds pay differentials and bonuses for STEM teachers, who theoretically could choose more lucrative careers with their given training—the assumptions of the alternative job model may not hold for special education teachers. Specifically, unlike teachers in math or science, special educators’ directly transferable skills may not apply to positions with higher salaries outside of education.

On the other hand, bonuses can also be used to offset the cost to teachers associated with more demanding jobs. Teaching students with disabilities can be challenging. Special educators have additional roles and responsibilities, beyond those of a general education teacher providing traditional instruction to students; in addition to their own instructional responsibilities, they are also expected to assist and support other teachers, coordinate related service providers to provide additional support to students, schedule meetings to plan and write individualized education plans, and implement progress monitoring and services required by these plans (Bettini et al., 2021). Further, they are often engaging in these responsibilities with limited access to the supports that are afforded to general educators, such as dedicated planning time and curricular resources (e.g., Billingsley & Bettini, 2019; Gesel et al., 2022; O’Brien et al., 2019). Indeed, some prior research indicates that they report more overwhelming workloads than their general education colleagues (Bettini et al., 2017). The results of a discrete choice survey experiment of teachers suggest that teachers are willing to forgo salary increases for additional supports with students with disabilities (Lovison & Mo, 2022).

As illustrated by Glazerman et al. (2013), teachers can be enticed to take potentially more challenging positions when the financial gain is sufficient. The null effects identified for special educators in the North Carolina bonus program may illustrate that the financial gains did not offset the additional and unique challenges to being a special education teacher. Notably, the

bonus was relatively small (\$1,800), whereas the bonus in the Glazerman et al. study was substantial, \$20,000 divided across 2 years. It is possible that special educators may require a larger bonus to remain in their position than teachers in other hard-to-staff areas.

Finally, another key difference between special education positions and other teaching positions in public schools is that, at least in some states, a large pool of teachers within the existing labor market possesses all the right credentials to teach special education but are choosing not to do so (e.g., Boe, 2006; Theobald et al., 2021). In fact, Theobald et al. (2021) show that one of the predominant sources of special education teacher shortages in Washington state is the movement of “dual-licensed” teachers—i.e., teachers who are credentialed to teach both special education and another subject—from special education to general education positions early in their teaching careers. Thus, one potential mechanism through which financial incentives could help address special education teacher shortages is by incentivizing these teachers to move back into special education teaching positions.

### **3. Policy Context**

Hawai‘i Public Schools have educated about 170,000 students (Hawai‘i Department of Education [HDOE], 2022a) and employed about 12,500 teachers (HDOE, 2022b) in recent school years, and have historically experienced pervasive teacher shortages that are particularly acute in special education due to a variety of factors including isolation, high rates of attrition, and few local teacher education programs (Kim, 2022). Importantly, Hawai‘i is the only state in the country with a single local education agency (i.e., school district), meaning that Hawai‘i Public Schools plays the roles of both district and state leadership. Within Hawai‘i Public Schools, schools are organized into complex areas containing two to four complexes of elementary, middle, and high schools. Another important difference between Hawai‘i and other school and district settings in the United States is the demographic makeup of the student body

and teaching staff; for example, about a quarter of teachers in Hawai‘i are White, another quarter are two or more races, another quarter are Japanese, and most of the other 25% are from a variety of Asian, Pacific, and Native Hawai‘ian backgrounds (author’s calculations). As Hawai‘i is a very unique setting in terms of the diversity of the teaching workforce, we do not investigate the impacts of the bonus policy on teacher workforce diversity explicitly, as the findings would be unlikely to generalize to other settings and the racial/ethnic categories in Hawai‘i are not easily categorized into sub-groups with sufficient sample sizes for appropriate heterogeneity analyses.

We now provide relevant background about the bonus policy that is the focus of this analysis. With support from Governor David Ige and the Hawai‘i State Board of Education (HBOE), the HDOE implemented differentiated pay for educators teaching in fall 2020 to increase compensation for classroom teachers in areas that experienced teacher shortages: special education, Hawai‘ian language immersion programs, and hard-to-staff locations. Importantly, the earliest written announcement of this policy was dated December 2019 (HDOE, 2019), which we argue is late enough not to have influenced staffing as of October 2019 (i.e., the beginning of the 2019–20 school year) but early enough to leave sufficient time for teachers to plan for their employment in October 2020 (i.e., the beginning of the 2020–21 school year). For the purposes of our statistical analysis, we therefore consider the first “treated” year to be 2020, as this is the first year that special education teachers were receiving bonuses. As documented by Kim (2022) and Lee (2021), although proposals were introduced in the 2021–22 school year to eliminate the bonus policy due to budget shortfalls, Senate Bill 2820 established these differentials permanently using state funds.

As described in the initial policy announcement (HDOE, 2019), pay differentials are unique to position type (i.e., special education or Hawai‘ian language immersion) and the

location of the teacher based on hard to staff designations: teachers of special education received a \$10,000 incentive, teachers teaching in Hawai‘ian language-immersion programs received an \$8,000 incentive and teachers in hard-to-staff school received incentives based on tier (Tier 1 \$3,000; Tier 2 \$5,000; Tier 3 \$7,500; Tier 4 \$8,000). Importantly, for the purposes of this analysis, our dataset does not identify Hawai‘ian language-immersion teachers. These non-special education teachers are included in the control group in our analysis even though they were receiving bonuses through this program. It is, therefore, a source of error. However, we do not view this as a huge limitation, as there are relatively few Hawai‘ian language immersion teachers in the state who received the bonus; as of January 2021, 94 Hawai‘ian language immersion teachers received the bonus, compared to 2,029 special education teachers (Lee, 2021). If anything, this biases the estimates of the impact of the bonus policy toward zero in our subsequent analysis.

Another important aspect of the policy is that incentives could be bundled so that a special education teacher in a hard-to-staff Tier 4 school would be eligible for the annual \$10,000 special education incentive as well as the annual \$8,000 incentive for working in a Tier 4 hard-to-staff school. For reference, a beginning teacher in this circumstance would earn a base salary of \$49,100 with the additional annual differential of \$18,000 for a total salary of \$67,100. Eligibility requirements included having a bachelor’s degree from an approved teacher preparation program and a license from the Hawai‘i Teacher Standards Board. Hard-to-staff schools are designated into tiers based on the number of met criteria: (1) complex areas with low rates of qualified teachers from state-approved teacher education programs over the last 3 years, (2) geographic isolation (25+ miles from an urban center), (3) complex areas with combined vacancy and emergency hire rates higher than 10% in school years 2016–17 and 2017–18 (SB



No. 2820, 2022). The more met criteria, the higher the tier number and the greater the pay differential for teachers teaching in these locations.

Finally, an extremely important piece of policy context not unique to Hawai‘i is the onset of the COVID-19 pandemic in March 2020—i.e., between the announcement of the bonus policy in December 2019 and the implementation of the policy in the 2020–21 school year—that closed schools for in-person instruction in Hawai‘i and across the country. Hawai‘i Public Schools were closed for in-person instruction through the end of the 2019–20 school year, and schools implemented a variety of modalities including in-person learning, distance learning, and hybrid learning throughout the 2020–21 school year (HBOE, 2020). The pandemic undoubtedly had a direct impact on teachers’ career decisions, as it did in other states that tended to see a decrease in teacher attrition after the first year of the pandemic and then steadily increasing attrition after the second and third pandemic years (e.g., Camp et al., 2023; Diliberti & Schwartz, 2023; Goldhaber & Theobald, 2023). But a key identifying assumption of this analysis is that the pandemic did not have a *disproportionate* impact on special education teachers relative to general education teachers, so that patterns of shortages for general education teachers can serve as an appropriate counterfactual for special education teachers after the introduction of the bonus policy. We are not aware of other large-scale pandemic-era interventions in Hawai‘i that targeted special education teachers, but given the empirical evidence on the plummeting of teachers’ perceptions of their working conditions during the pandemic (e.g., Kraft et al., 2021) and the importance of working conditions for special education teachers in particular (e.g., Billingsley et al., 2020), we argue that, if anything, the pandemic likely had a disproportionately negative impact on special education teachers, which would again bias our estimates toward zero.

## **4. Data and Analytic Approach**

### **4.1 Data**

The educator position records used for this study were provided to the research team by HDOE in April 2023 and include every position in the Hawai‘i public school system between 2014–15 and 2022–23, linked to the specific school, school tier designation (i.e., Tier 1 through Tier 4 as described above), and whether the position is in special education. For positions that are not vacant, each observation also includes data about the individual filling the position, including their race/ethnicity, an indicator for whether the teacher is licensed by the Hawai‘i Teacher Standards Board, and an anonymous personnel ID that allows us to track individual employees across years. For this study, we exclude central office vacancies at the state complex area, and complex levels (3.68% of all records) and positions in community schools for adults (0.07% of all records). The analytic sample includes seven districts, 15 complex areas, and 261 schools with a sample size of more than 115,000 positions across 9 school years.

We create four outcome measures for this study. For RQ1 and RQ3, we first create binary indicators for whether each position in each year is vacant. Figure 1 summarizes the proportion of special education and not special education positions that were vacant in October of each school year. In each year between 2014 and 2019 (i.e., the year before the bonus policy), 5%–6% of special education positions were vacant compared to 1%–3% of not special education positions. But the proportion of vacant special education positions decreased to less than 4% in 2020, the first year of the bonus policy, and although this proportion increased in 2021 and 2022 (up to more than 6% in the most recent 2022–23 school year) the increases were similar to those seen in general education. As formalized below, the intuition behind our analysis is that the trend in vacancies for not special education teaching positions in Figure 1 provides an appropriate

counterfactual after the introduction of the bonus policy for the trend in vacancies for special education teaching positions.

The second outcome of our analysis of RQs 1 and 3 expands the definition of “shortage” to include positions filled by an unlicensed teacher; i.e., the outcome is an indicator for whether a position is vacant *or* filled by an unlicensed teacher. We summarize the proportion of positions that fall into this category by position type and in year in Figure 2. The proportion of special education positions that were vacant or filled by an unlicensed teacher ranged from 12%–17% prior to the introduction of the bonus policy—compared to 4%–7% of general education positions—but decreased to close to 10% in the first 2 years after the policy and was still 13% in the most recent 2022–23 school year. As with vacancies, the proportion of general education positions that were vacant or filled by an unlicensed teacher also increased in recent years to a high of 8% in 2022–23.

Our analysis of RQ3 investigates heterogeneity in these outcomes across school tiers in which all teachers received additional bonuses. We therefore provide summary statistics analogous to Figures 1 and 2 separately by school tier (Tier 0 = no bonus) in Figures 3 and 4. There are three important patterns in these figures. First, partially by definition as school tiers were selected in part because of prior rates of vacancies, schools in Tiers 3 and 4 schools particularly have higher baseline rates of positions that are vacant and/or filled by unlicensed teachers; for example, more than 10% of all special education positions were vacant in Tier 4 schools in every year from 2016 to 2019, whereas more than 25% of special education positions either were vacant or filled by an unlicensed teacher in these schools in every year from 2014 to 2019. Second, estimates in Tier 0 schools are more precise and track overall patterns in Figures 1 and 2 closely because more than 80% (81.5%) of position-year observations are in these schools,

compared to 2.8% in Tier 1, 6.8% in Tier 2, 1.7% in Tier 3, and 7.1% in Tier 4. Finally, there are notable drops in the proportion of positions that are vacant and/or filled by unlicensed teachers in Tier 4 schools particularly after the introduction of the bonus policy in 2020; specifically, the proportion of vacancies dropped from more than 10% in 2016–2019 to less than 5% in 2020–2022 in Tier 4 schools, whereas the proportion of positions that are vacant or filled by unlicensed teachers dropped from more than 25% in 2014–2019 to slightly more than 10% in 2020–2022.

We now summarize the two outcomes for our analysis of RQ2. The first outcome follows prior literature of teacher attrition and considers binary indicators for whether each teacher in the state left the state’s teaching workforce the following year (i.e., is not observed as a teacher in the following school year). Note that, although previous studies of teacher turnover also consider teacher mobility between schools (e.g., Gilmour & Wehby, 2020), we do not explicitly consider between-school mobility as an outcome in this study because between-school mobility does not impact either of the primary state-level outcomes we consider as part of RQ1.

Figure 5 summarizes overall attrition rates by position type and school year; in Figure 5 and subsequent attrition analyses, the attrition rate for 2014 is the proportion of teachers in October 2014 (i.e., the 2014–15 school year) who did not return to teach in October 2015 (i.e., the 2015–16 school year). Attrition rates for special education teachers are 12%–15% in each year of data and are consistently 3 to 4 percentage points higher than the analogous attrition rates for general education teachers. Unlike the prior two outcomes, these patterns do not change substantially after the pandemic, though we will formalize this comparison in the analysis described below.

Finally, to create a measure that will allow us to consider whether “open positions” were filled by an existing teacher for the second half of RQ2, we first need to categorize different

teaching positions in each year and define what we mean by open positions given that we do not observe hiring data. In Figure 6, we show that all teaching positions in each year and position type can be placed into one of four categories: filled by an existing teacher from the same position type (e.g., a special education teacher who also was a special education teacher last year); filled by an existing teacher from the other position type (e.g., a special education teacher who was a general education teacher last year); filled by a new teacher (i.e., a teacher who was not in the workforce last year); or a vacant position. Figure 6 shows that the second category increases substantially for special education positions after the introduction of the policy.

But to formalize this as a regression outcome, we define “open positions” as those that fall into one of the last three categories; these are positions that are not filled by a teacher who already was in the same position type last year. We then calculate the proportion of these positions that are filled by an existing teacher from the other position type and plot the proportion of these positions by type and year in Figure 7. Although 9%–14% of open special education positions were filled by an existing general education teacher before the introduction of the bonus policy, Figure 7 shows this percentage increased to 27% in the first year of the policy (i.e., a quarter of all open special education positions were filled by an existing general education teacher in 2020). We consider the binary indicators summarized in Figure 7 as the final outcome in our regression analysis associated with RQ2, as it provides descriptive insight into the extent to which hiring patterns changed after the introduction of the bonus policy.

#### **4.2 Analytic Approach**

Our analytic approach leverages the fact that there was a one-time policy change that impacted only a subset of positions in the state (RQs 1 and 2) and may have impacted these positions differently depending on the school setting of the position (RQ3). Specifically, for

RQ1, we follow Feng and Sass (2018) and estimate difference-in-differences (DID) models predicting the outcomes associated with each research question. For RQ1, let  $Y_{prst}$  be an indicator for whether position  $p$  that is in instructional role  $r$  (special or general educator), school  $s$ , and year  $t$  is either vacant or filled with an unlicensed teacher. We model some function of the probability that this indicator is 1,  $p_{prst} = \Pr(Y_{prst} = 1)$ —either the log odds with a logistic regression or the probability itself with OLS—as a function of an indicator for years in which special education teachers were receiving a \$10,000 bonus,  $I_{t \geq 2020}$ , an indicator for instructional role  $r$ ,  $I_r$ , and the interaction between these two indicators:

$$f(p_{prst}) = \alpha_0 + \alpha_1 I_{t \geq 2020} + \alpha_2 I_r + \alpha_3 I_r * I_{t \geq 2020} + \alpha_s + \varepsilon_{prst} \quad (1)$$

The coefficient of interest,  $\alpha_3$ , can be interpreted as the effect of the special education bonus on the probability that positions are either vacant or filled by an unlicensed teacher. We estimate some specifications without a school fixed effect  $\alpha_s$ , but our preferred specification includes these fixed effects to make comparisons within the same schools over time. Finally, in some specifications, we also include school-by-role fixed effects  $\alpha_{sr}$  and school-by-year fixed effects  $\alpha_{st}$  ensure that comparisons are made within specific school roles or school years. We cluster all standard errors at the school level to account for correlations between measures of shortage within the same school across years. Because the treatment in Equation 1 is in a single year, this approach is not subject to issues raised with two-way fixed effects DID models that rely on identification due to staggered adoption (Roth et al., 2023), but it is subject to the “parallel trends” assumption that outcomes for general education positions provide an appropriate counterfactual for special education positions after the introduction of the policy. We therefore test for parallel trends in the 6 years before the introduction of the policy in the 2020

school year. Specifically, we estimate an event study model that replaces the posttreatment indicator in Equation 1 with a series of year indicators  $I_t$ :

$$f(p_{prst}) = \beta_0 + \beta_{1t}I_t + \beta_2I_r + \beta_{3t}I_r * I_t + \beta_{sr} + \beta_{st} + \varepsilon_{lrst} \quad (2)$$

As is conventional, the year immediately before the incentives took effect (2019) is the reference period and omitted. Precisely estimated zero  $\beta_{3t}$  for years before the policy took effect would provide evidence that the parallel trends assumption cannot be rejected. For  $t \geq 2020$ , estimates for  $\beta_{3t}$  allow us to explore how the effect of incentives may have evolved over time. Even though our research design does not involve staggered implementation timing, Sun and Abraham (2021) demonstrate that the traditional event study estimator for  $\beta_{3t}$  is a linear combination of group-specific effects from all relative periods. These estimates are biased in the presence of heterogeneous treatment effects. To examine the sensitivity of our findings to treatment effect heterogeneity, we follow Callaway and Sant’Anna (2021) and estimate dynamic treatment effects using their doubly robust estimator.

The first part of RQ2 can be addressed by replacing the dependent variable in equations (1) and (2) with measures of teacher attrition summarized in Figure 5. Specifically, we can lag the outcome by one year and estimate OLS and logit specifications of equations 1 and 2 in which the outcome is whether a teacher in a given position leaves the workforce the following year. And then for the second part of RQ2, we limit the sample to “open positions” and substitute the outcome summarized in Figure 7—that is, whether each open position is filled by an existing teacher from a different position type. Importantly, we do not view this subanalysis as causal, as the bonus policy likely impacted both switches from general education to special education and vice versa (i.e., general education positions do not necessarily provide a proper “counterfactual”

for trends in the absence of this policy). But we still view this as a useful exercise for formalizing the trends documented in Figure 7.

Finally, RQ3 can be addressed by estimating equations (1) and (2) separately for schools in the four tiers that are eligible for hard-to-staff bonuses and comparing the estimated effect with similar effects for schools in not hard-to-staff schools (RQ3). Our preferred DID model (i.e., Equation 1) specifies this as a triple difference model that fully interacts position type, school tier, and the post policy indicator; in this model, the three-way interaction terms provide information about how the impact of the policy varied within schools designated into different tiers for the purposes of the school bonus policy.

## **5. Results**

### ***5.1 What is the average causal effect of the \$10,000 bonus policy on special education teacher shortages in Hawai‘i, as measured by the proportion of positions that are vacant and/or by the proportion of positions filled by unlicensed teachers?***

In Table 1, we report estimates from the DID model in Equation 1. The coefficient of interest, “Special Education \* Post Bonus”, can be interpreted as the causal effect of the special education bonus on the probability that a position is vacant (Panel A) or vacant or filled by an unlicensed teacher (Panel B) under the identifying assumptions discussed in Section 4. We first report a naïve specification that omits school fixed effects (column 1), and then report estimates from our preferred specification with school fixed effects first as odds ratios (column 2) and then as average marginal effects from this logistic regression (column 3). To explore the sensitivity of our findings to the specification of the DID model, we then report the same specification estimated as an OLS model in column 4, and then add school-by-role fixed effects (column 5) and school-by-year fixed effects (column 6) to the model to make comparisons between increasingly narrow groups of positions.



Focusing first on Panel A, the odds ratio estimate in our preferred specification in column 2 (0.68) implies that the introduction of the special education bonus predicts a 32% decrease in the proportion of special education teacher positions in the state that were vacant. This estimate is similar whether or not we include school fixed effects (columns 1 and 2), and expressed as a marginal effect equates to a 1.2 percentage point reduction in special education vacancies, whether or not we estimate this as the average marginal effect calculated after logistic regression (column 3) or from a linear probability OLS model (column 4). Finally, this marginal effect is very similar in models that explicitly make comparisons within schools and roles (column 5) and within schools and years (column 6).

Interpreting these estimates as causal relies on a number of important identifying assumptions, the most important of which is that trends in general education teacher vacancies after the introduction of the bonus policy can serve as an appropriate counterfactual for trends in special education teacher vacancies in the absence of the bonus policy. The typical approach to test this is by estimating an event study model (Equation 2) and testing for “parallel trends” in the six years prior to the introduction of the policy in the 2020 school year. Figure 8 shows the estimates from the event study model version of our preferred model specification (logistic regression). There are two important conclusions from Figure 8. First, there is no visual evidence of parallel trends in vacancies between the general education and special education positions, which is bolstered by formal tests that fail to reject the null hypothesis of parallel trends ( $p = 0.65$  for the logistic regression model and  $p = 0.91$  for the Callaway and Sant’Anna [2021] estimator). Second, the estimates from the three treatment years in Figure 8 show relatively consistent effects across years.

We now turn to Panel B of Table 1, in which the outcome is whether a given position is either vacant or filled by an unlicensed teacher. The odds ratio in column 2 implies that the introduction of the bonus policy predicts a 35.5% reduction (or 4 percentage point reduction under either method of calculating marginal effects in columns 3 and 4) in this proportion for special education positions. Again, these estimates are extremely robust to different model specifications in columns 5 and 6, and the event study estimates in Figure 9 show no significant pre-trends ( $p = 0.49$  for the logistic regression model and  $p = 0.30$  for the Callaway and Sant’Anna [2021] estimator). However, unlike the event study for vacancies, the estimates for post-policy treatment effects suggest that the impact of the policy increased with time, meaning that the reduction in the proportion of positions that were vacant or filled by an unlicensed teacher only increased in the second and third years of the policy.

## ***5.2 To what extent are these effects driven by changes in patterns of teacher attrition and movement between special education and general education teaching positions?***

We now turn to our investigations of potential mechanisms for the effects documented in Section 5.1. Table 2 follows the same format as Table 1 except the outcome in Panel A is an indicator for whether each teacher in the state leaves the state’s teaching workforce at the end of the year, whereas the models in Panel B are limited to open positions as defined in Section 4 and the outcome is an indicator for whether the open position is filled by an existing teacher from the other position type (general education or special education). In Panel A, we find no significant relationship between the introduction of the bonus policy and the attrition rates of special education teachers. For example, the magnitude of the marginal effect from our preferred specification in column 3 (-.002) and the associated standard error (.005) imply that we can rule out with 95% confidence retention effects of greater than about 1 percentage point in either

direction. This is borne out visually from the event study figure in Figure 10, which shows neither a significant pre- or post-trend in the attrition rate of special education teachers relative to general education teachers. Thus, we do not have sufficient evidence to conclude that the special education bonus policy had a significant impact on special education teacher attrition.

But in Panel B—and as previewed by the summary statistics in Figure 7—we estimate extremely large relationships between the introduction of the policy and the probability that an open special education position was filled by an existing general education teacher; i.e., the estimate in column 2 implies that this probability more than doubled (108% larger) with an increase of about 8 percentage points regardless of specification (columns 3–6). The event study figure in Figure 11 shows no significant pre-trend, and that this increase was disproportionately driven by a large spike in the movement of general education teachers into open special education positions in the first year of the bonus policy. Although we view these estimates as descriptive due to the identification issues discussed in Section 4, our interpretation of Table 2 is that the impacts of the bonus policy on special education teacher shortages were driven almost entirely by the movement of general education teachers into special education teaching positions as opposed to the retention of existing special education teachers in the workforce.

### ***5.3 How does the effect of the \$10,000 bonus policy vary between hard-to-staff schools in which all teachers received additional bonuses and other schools?***

Finally, we report estimates from the DDD model described in Section 4 that allows us to investigate heterogeneity in the relationships discussed in Section 5.1—both impacts on vacancies (columns 1–3) and vacancies plus unlicensed teachers (columns 4–6)—by the various tiers of schools in which teachers received schoolwide bonuses in addition to subject-specific bonuses in special education. The first row of Table 3 can be interpreted as the effect of the

introduction of the policy in Tier 0 schools (i.e., schools in which there were no schoolwide bonuses), whereas the interaction terms with the four school tiers show how this effect varies by school tier.

Focusing first on vacancies (columns 1–3), we do not find that the bonus policy had a significant impact on the proportion of special education vacancies in Tier 0 schools. Instead, the negative impact is disproportionately driven by a large negative impact (7.2 percentage points combining the main and interaction effects) in Tier 4 schools. As shown in the separate event study figures in Figure 12, the impact of the bonus policy in Tier 4 schools was particularly large in the most recent (2022–23) school year, with a reduction in vacancies of nearly 10 percentage points in these schools.

On the other hand, the bonus policy did have a significant negative impact on the proportion of positions that were vacant or filled by an unlicensed teacher in Tier 0 schools (2.4 percentage points, columns 4–6), but again, the effects are larger in Tier 4 and (in some specifications) Tier 2 and Tier 3 schools. Put together, the estimates imply that the introduction of the bonus policy reduced the proportion of positions that were vacant or filled by an unlicensed teacher in Tier 4 schools by more than 15 percentage points. And as shown in the event study figure for these regressions in Figure 13, this effect was particularly large in the most recent 2022–23 school year, whether the effect size was more than 20 percentage points relative to the pretreatment year.

## **6. Discussion**

We found that Hawai‘i’s bonus program had an immediate, significant, and meaningful impact, substantially reducing the proportion of special education teaching positions that were vacant or filled by unlicensed personnel. These results indicate that substantial financial bonuses

can meaningfully reduce special education teacher shortages and thereby increase the likelihood that students with disabilities have qualified special education teachers.

Importantly, however, the bonus policy had no effect on retention among teachers who already were serving as special educators. Rather, the effect was driven by internal transfers—that is, movement of personnel who were not previously teaching special education into special education teaching positions. This indicates that the policy was insufficient to stem attrition of current special education teachers, but it was sufficient to induce personnel who had not previously been teaching special education to change positions. This could be because moving from general education to special education teaching positions does not require teachers to change schools; they can make this switch without leaving the colleagues and community with whom they already have relationships, and without moving. Either way, the policy worked by redistributing teachers within Hawai‘i’s education system, yielding a more equitable distribution of current personnel. This echoes recent findings from Chile (Elacqua et al., 2022) and sources of special education teacher shortages in Washington State (Theobald et al., 2021) and has broad implications for efforts to improve equity in the assignment of teachers to students with disabilities in public schools (e.g., Lai et al., 2021); i.e., enticing general education teachers with dual licenses to move into special education positions has long been a potential solution to special education teacher shortages that has been “hiding in plain sight” (DeArmond, 2023; Theobald et al., 2021).

Yet, the fact that the bonus policy did not impact retention of current special education teachers also raises questions about how enduring the policy’s effects will be. Among teachers who switched positions, will the incentive be sufficient to induce them to stay for the long term? Perhaps teachers who transferred are individuals who are especially responsive to financial

incentives and continued bonuses will be sufficient to retain them, even though the bonus was not sufficient to retain the special educators already in these positions. If this is the case, the policy will likely have long-term positive impacts on special education teacher shortages.

Alternatively, perhaps once these educators are teaching special education, they may determine that the bonus is insufficient given the demands of the job, and thus they may choose to leave or switch back to other teaching positions. In that case, the policy will have provided a short-term solution to an urgent crisis, but other solutions (e.g., improving working conditions) may be needed to ensure a robust special education teacher workforce in the long term. Understanding how the bonus impacts long-term retention of personnel who switched positions will be crucial for evaluating the overall effects of financial incentive policies on special education teacher shortages.

Our null findings for retention also raise questions about what magnitude of financial incentives would be required to retain current special education teachers. Examining Florida's financial incentive policies from the late 1990s and early 2000s, Feng and Sass (2017) found that bonuses worth 5.7% of average teachers' salaries were sufficient to significantly reduce special education teacher retention by 10%–12%, whereas bonuses worth 2% of average salaries and loan forgiveness programs worth 3.4% of average salaries had no effect. The Hawai'i bonus of \$10,000 is worth more than 20% of the average beginning special education teacher's salary, representing a substantial bonus—far more than the bonuses that Feng and Sass (2017) found had an effect in Florida. On the other hand, Glazerman et al. (2013) found that a \$20,000 bonus was a sufficient transfer incentive to disadvantaged schools, so it is possible that even a \$10,000 bonus is insufficient to achieve desired retention effects.

Potentially, other factors could reduce the value of the \$10,000 bonus to Hawai‘i’s special education teachers. For example, if special educators’ working conditions are especially challenging in Hawai‘i, the bonus could be insufficient to compensate for dealing with these conditions on a daily basis. Future investigations should consider incorporating other factors into the analysis, to consider how the effects of financial incentives may vary based on schools’ working conditions and other economic conditions (e.g., cost of living, current local job market); perhaps financial incentives may need to vary in ways that are calibrated to these other conditions.

We also did not examine heterogeneous effects of the policy based on teacher characteristics, and future research is needed to consider which teachers might be especially responsive to financial incentive policies. For example, prior research consistently indicates that teachers of color experience more substantial financial barriers to teaching; generations of racist systems have left teachers of color with less family financial wealth, such that they are less able to manage financial demands of teaching (e.g., the expectation to purchase classroom materials with personal funds; Scott, 2020). Thus, it is possible that the policy could have been especially impactful for teachers of color and teachers from lower income backgrounds. If this is the case, financial incentive policies could be especially useful for diversifying the special education teacher workforce. The policy also could have had disproportionate effects on general education teachers who already were dual licensed, as they could more rapidly transfer into special education. Given evidence that dual-licensed general educators are especially effective at teaching students with disabilities in general education classes (Gilmour, 2020; Goldman & Gilmour, 2021), this would raise concerns about potential iatrogenic effects of the policy on effective inclusion of students with disabilities. Future research on how the policy impacted

student outcomes among students with disabilities would be useful for understanding these potential effects, particularly if results were disaggregated by the setting in which students are served (i.e., general education or separate settings).

We also did not examine the effects of the bonus policy on preparation or entry into special education, as preparation data were unavailable in the dataset. But it is possible that bonuses may also have impacted the likelihood that prospective teachers would choose special education. To fully understand the bonus policy's impacts, future research should also examine the effects of the policy on enrollment in and completion of special education teacher preparation programs, as previewed in early work by Kim (2022).

Finally, we found that the effects of the special education teacher bonus were magnified in the hardest-to-staff schools. Because all teachers in these schools received bonuses, we cannot disentangle the extent to which this impact is due to the larger bonuses for special education teachers in these schools (up to \$18,000 in Tier 4 schools), how much is related to the higher baseline rates of vacancies and positions filled by unlicensed teachers in these schools (i.e., there was greater scope for change in these schools), and how much is related to the specific school and community contexts in these schools. But this finding has clear equity implications as it suggests that students with disabilities in historically hard-to-staff schools can be disproportionately benefitted by a bonus policy like this.

### ***Implications for Policymakers***

Our results have several key implications for policymakers. Foremost, substantial financial incentives can support a more equitable distribution of well-qualified teachers to students with disabilities, with particularly strong effects for special education teachers in hard-



to-staff schools. Several factors may have contributed to these strong positive effects, which policymakers interested in financial incentive policies should attempt to replicate.

First, the special education teacher bonus policy was applied across the board, to all teachers in the state. The simplicity of the policy may have made it easier to market to prospective beneficiaries. Further, the state had a dedicated marketing program, designed to ensure that current and prospective teachers understood the policy when making career decisions. Clear, consistent, and straightforward communication likely contributed to the strong positive effects of the policy.

Second, teachers did not have to navigate complicated bureaucratic processes in order to benefit, reducing the likelihood that intended beneficiaries did not get access to the bonus. These features likely eased implementation, eliminating the kinds of challenges that have limited the effects of prior financial incentive policies (e.g., North Carolina's bonus program from the late 1990s and early 2000s; Clotfelter et al., 2005, 2008). Attempts to replicate the policy should attend to Hawai'i's implementation strategy, drawing on the strengths of this approach to ensure the effectiveness of other financial incentive policies.

### ***Conclusion***

We found that Hawai'i's \$10,000 bonus policy had an immediate and significant impact on the likelihood that special education positions were filled by licensed personnel. The policy worked by inducing current teachers who were not in special education teaching positions to transfer into special education, especially in the hardest-to-staff schools, but it had no effect on the likelihood that current special education teachers remained in their positions. These findings raise many questions for future research about the effects of financial incentive policies on the teacher workforce. However, they do indicate that financial incentives may be a useful strategy

for policymakers interested in reducing inequities, based on students' identification with a disability, in access to well-qualified teachers.

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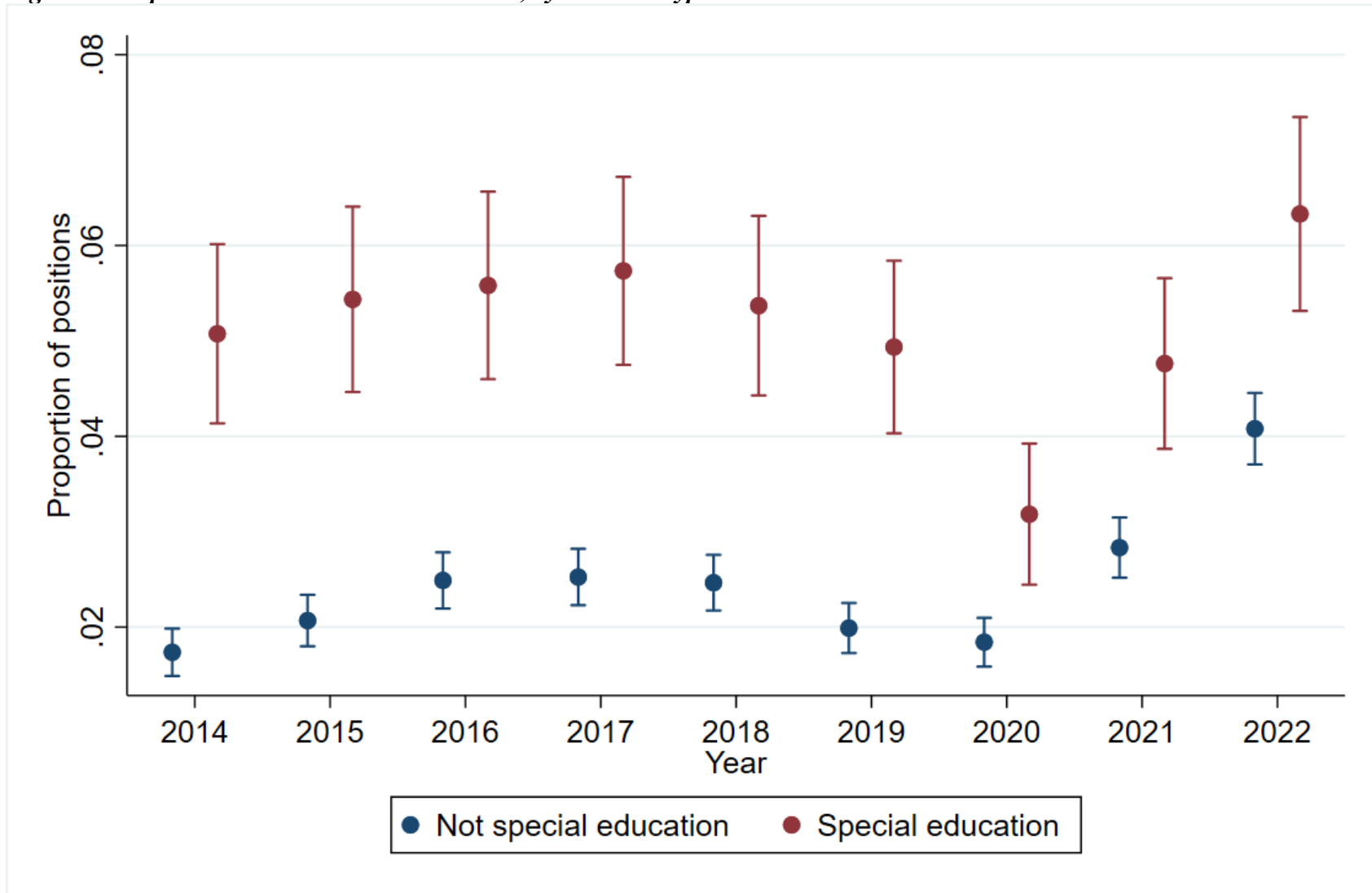
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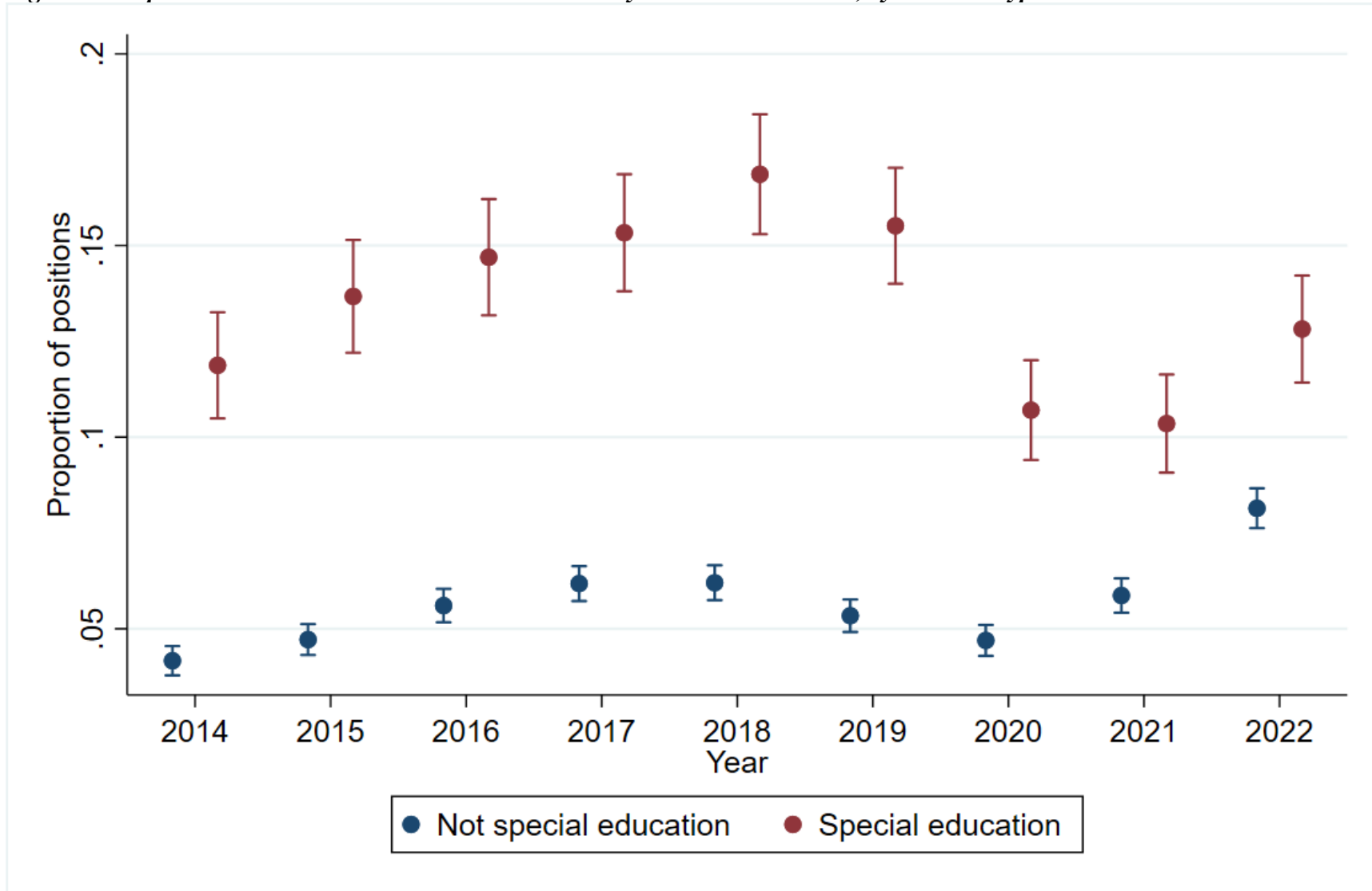
Figures and Tables

Figure 1. Proportion Positions That Are Vacant, by Position Type and Year



Note. 2014 refers to October 2014 of 2014–15 school year.

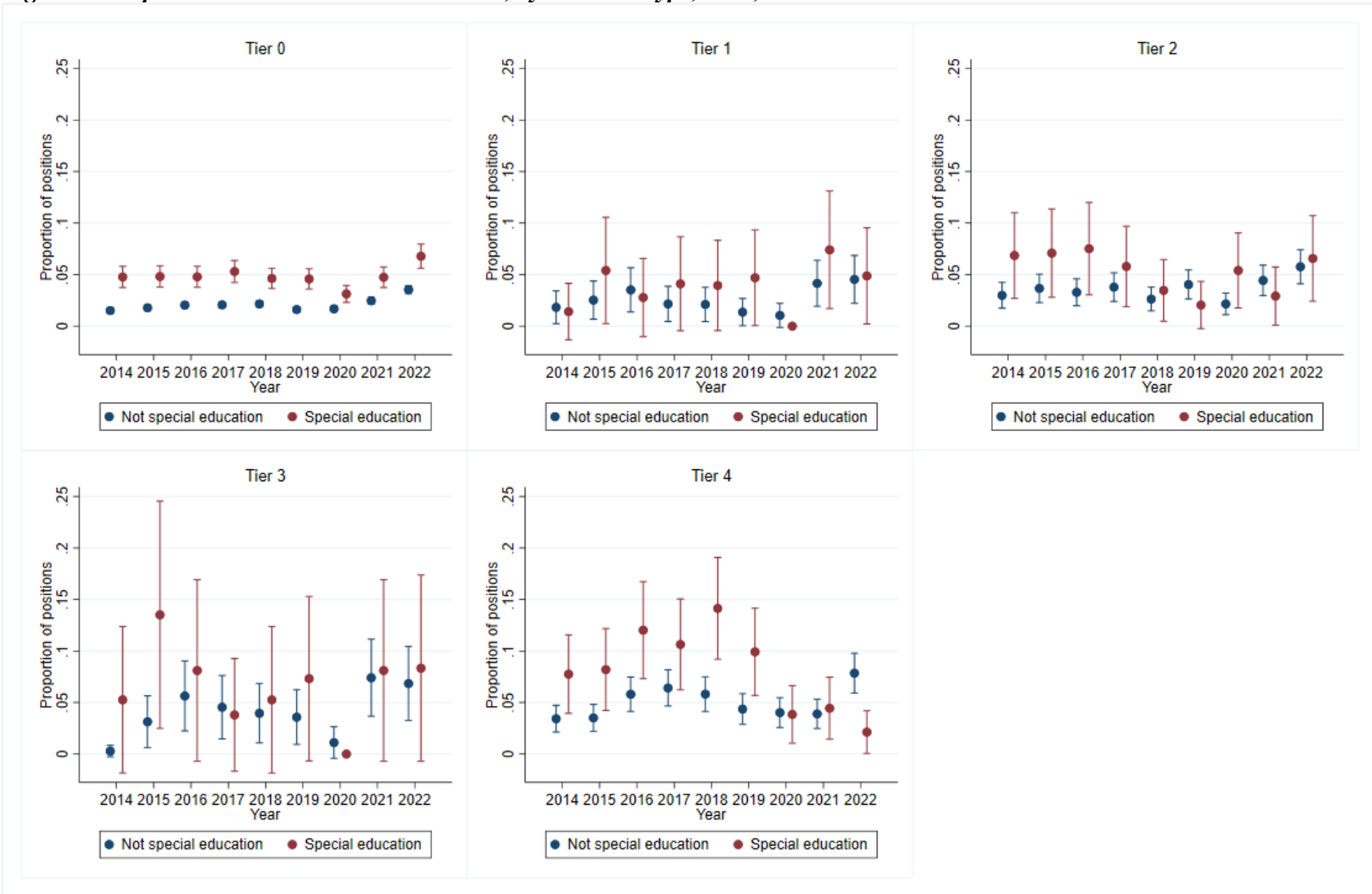
Figure 2. Proportion Positions That Are Vacant or Filled by Unlicensed Teacher, by Position Type and Year



Note. 2014 refers to October 2014 of 2014–15 school year.

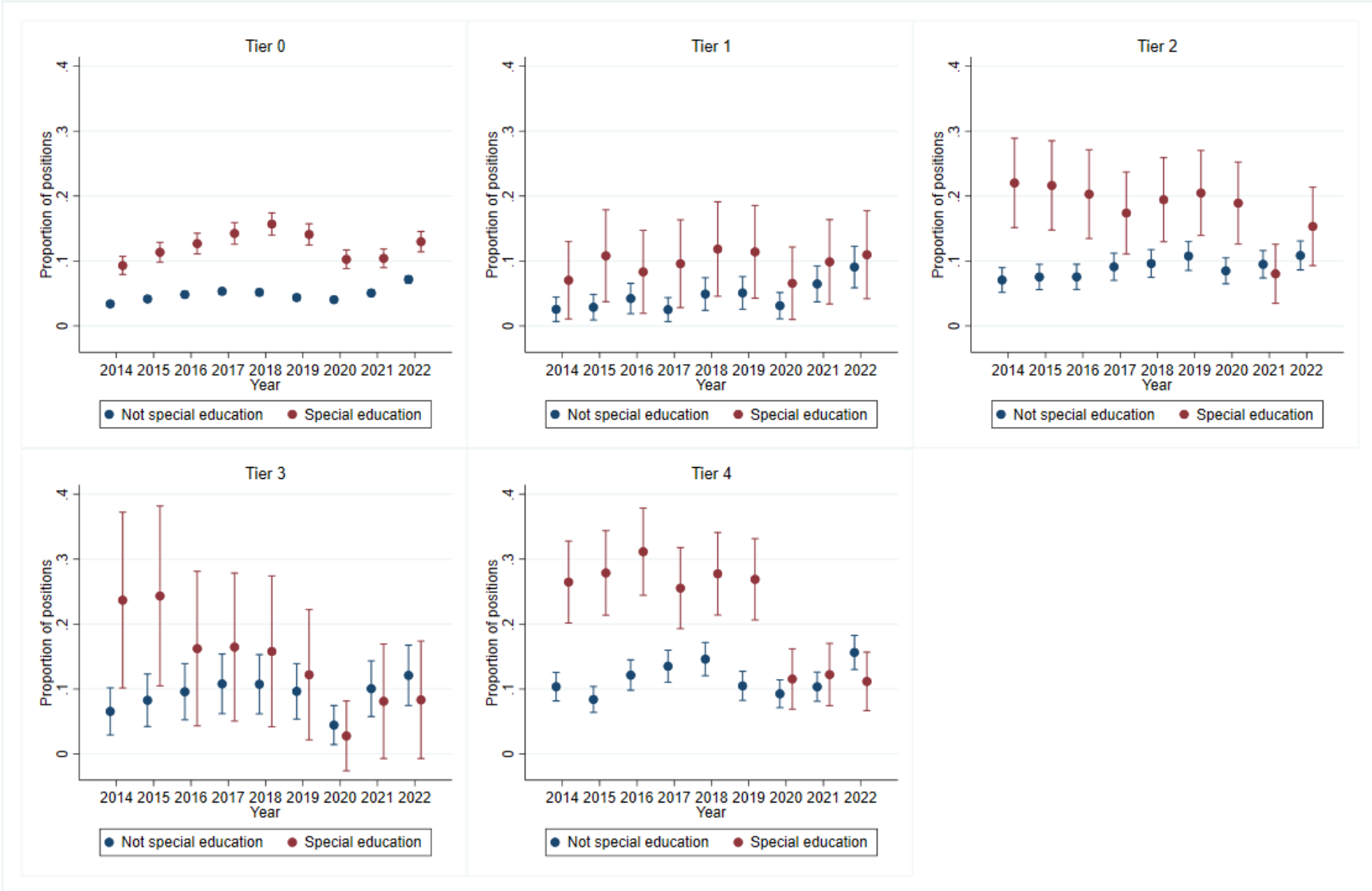


**Figure 3. Proportion Positions That Are Vacant, by Position Type, Year, and School Tier**



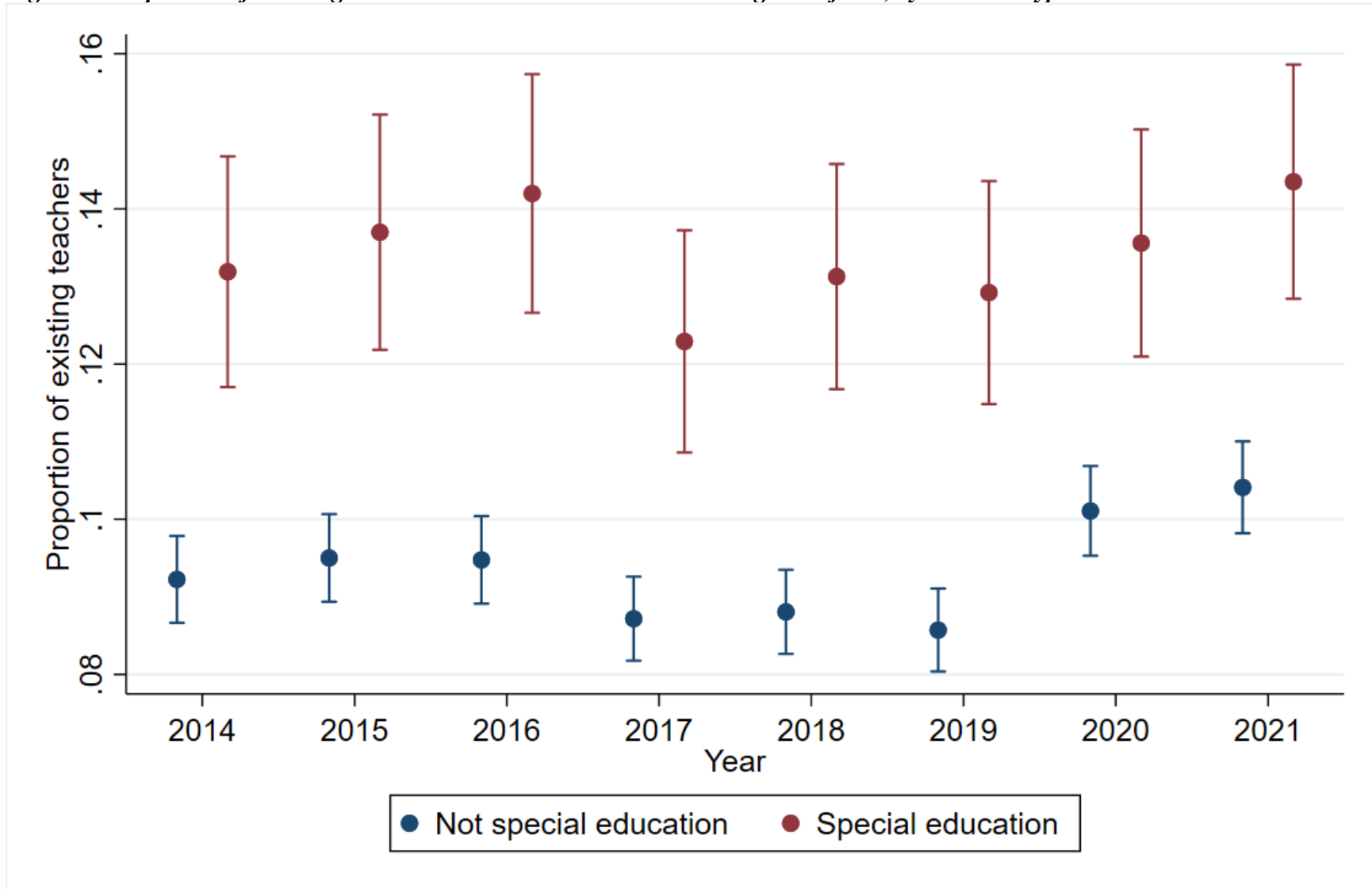
Note. 2014 refers to October 2014 of 2014–15 school year. SPED = special education position.

**Figure 4. Proportion Positions That Are Vacant or Filled by Unlicensed Teacher, by Position Type, Year, and School Tier**



Note. 2014 refers to October 2014 of 2014–15 school year. SPED = special education position.

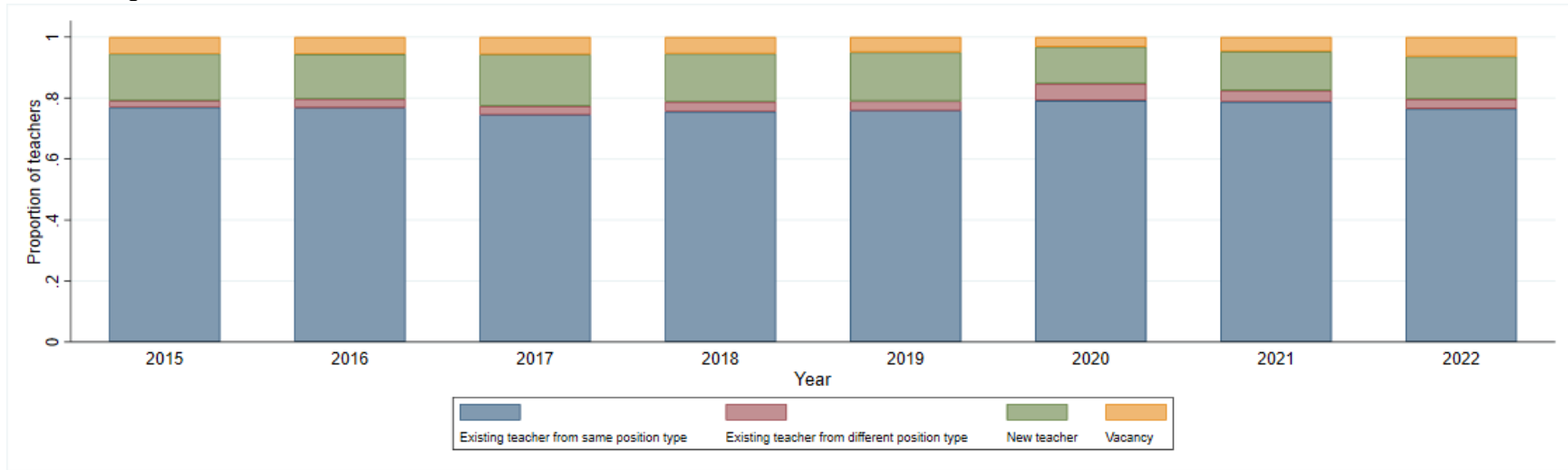
Figure 5. Proportion of Existing Teachers who Leave the State Teaching Workforce, by Position Type and Year



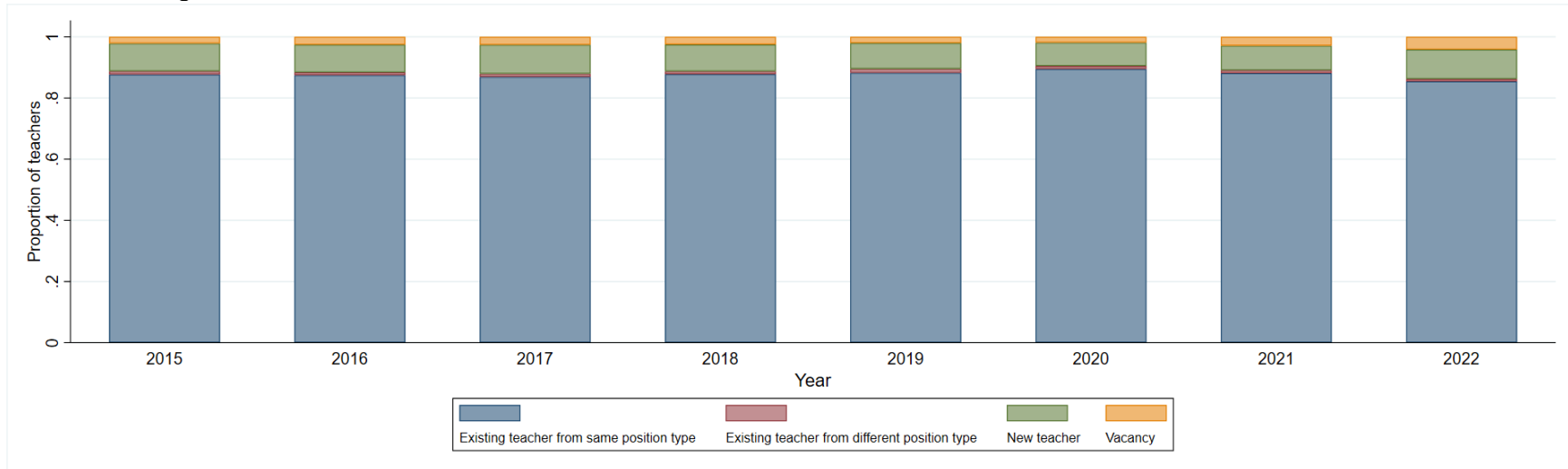
Note. 2014 refers to October 2014 of 2014–15 school year.

**Figure 6. Position Categorization, by Position Type and Year**

**Panel A. Special Education Positions**

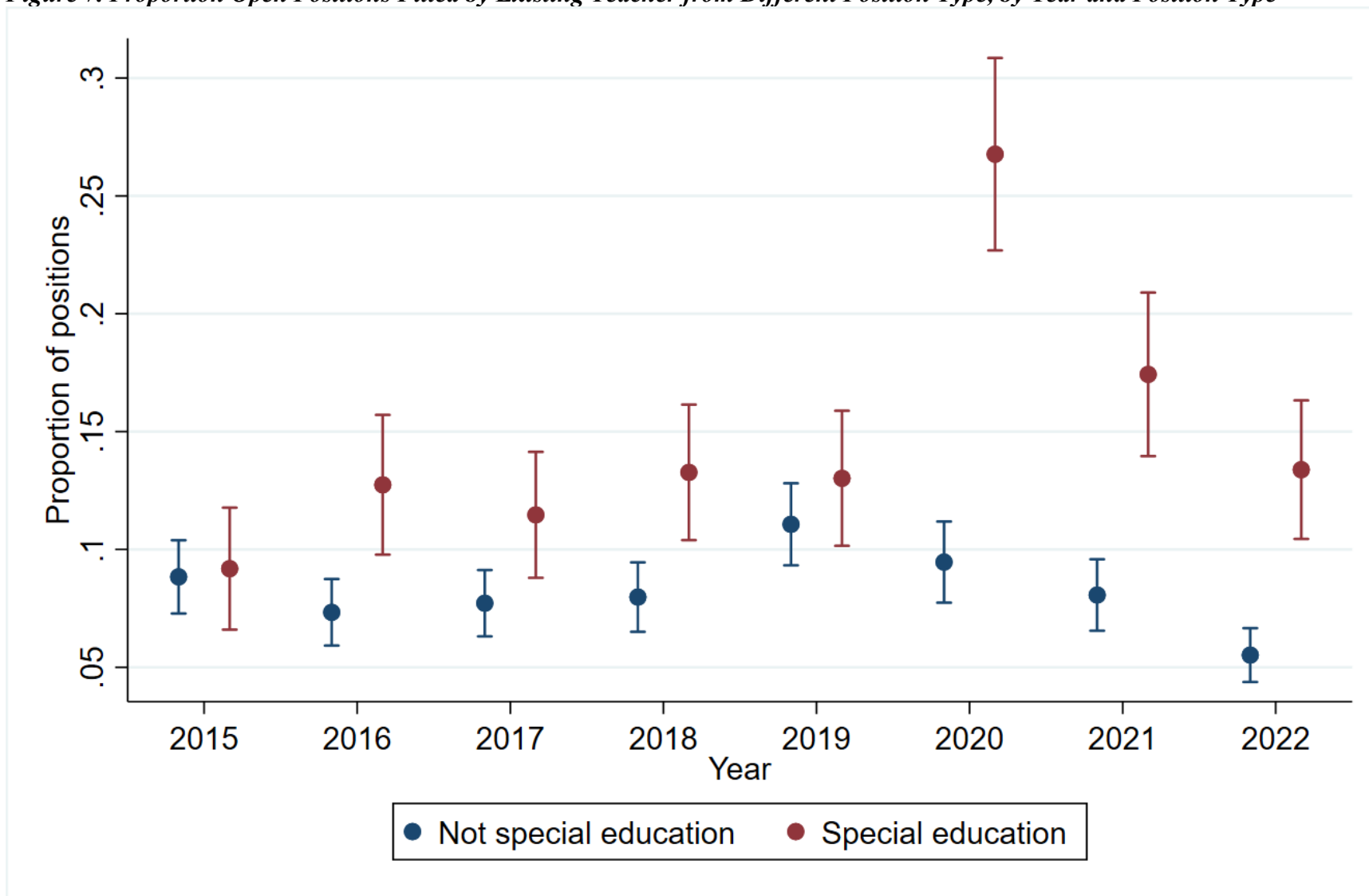


**Panel B. Not Special Education Positions**



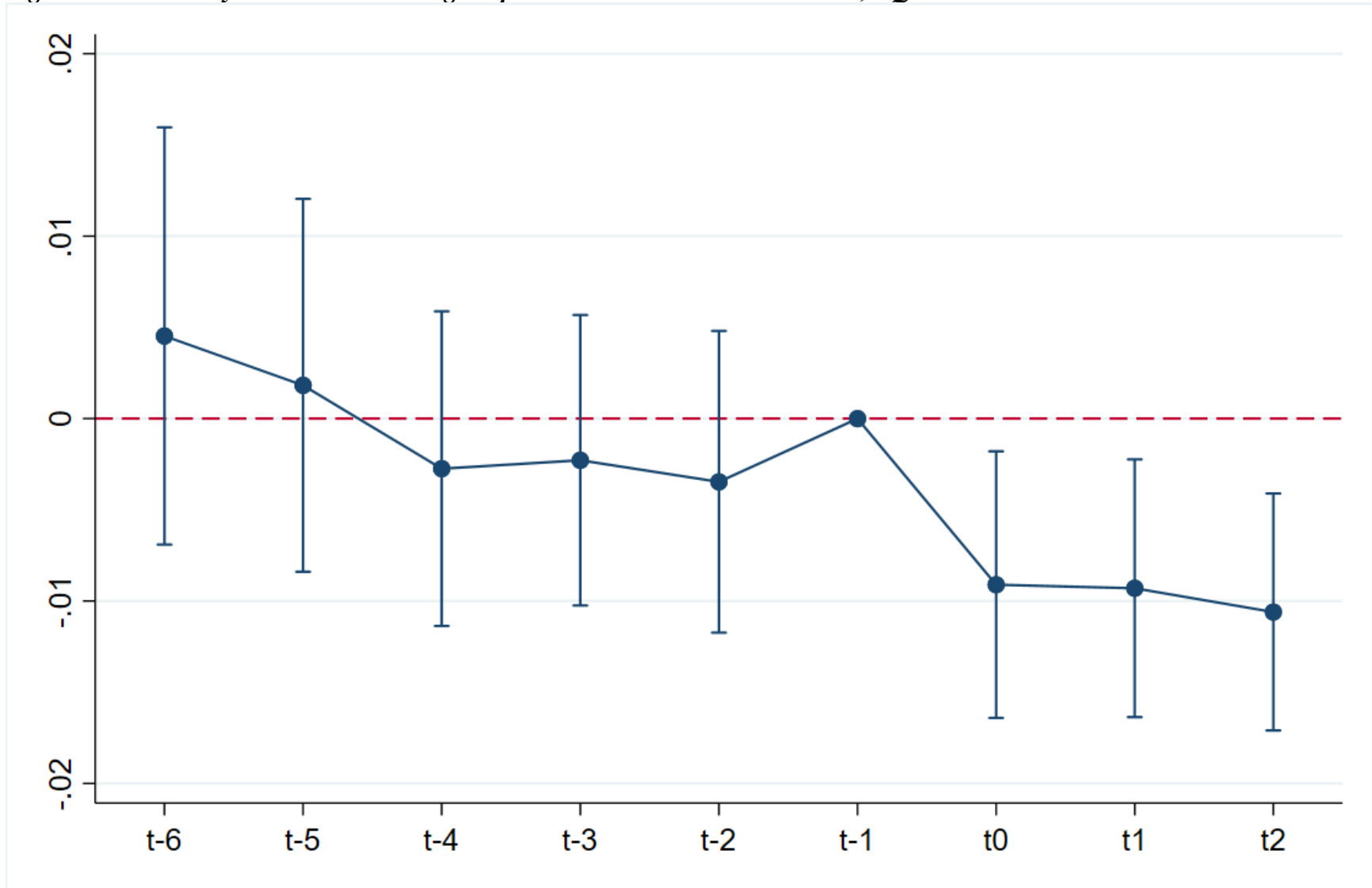
*Note.* 2015 refers to October 2015 of 2015–16 school year.

Figure 7. Proportion Open Positions Filled by Existing Teacher from Different Position Type, by Year and Position Type



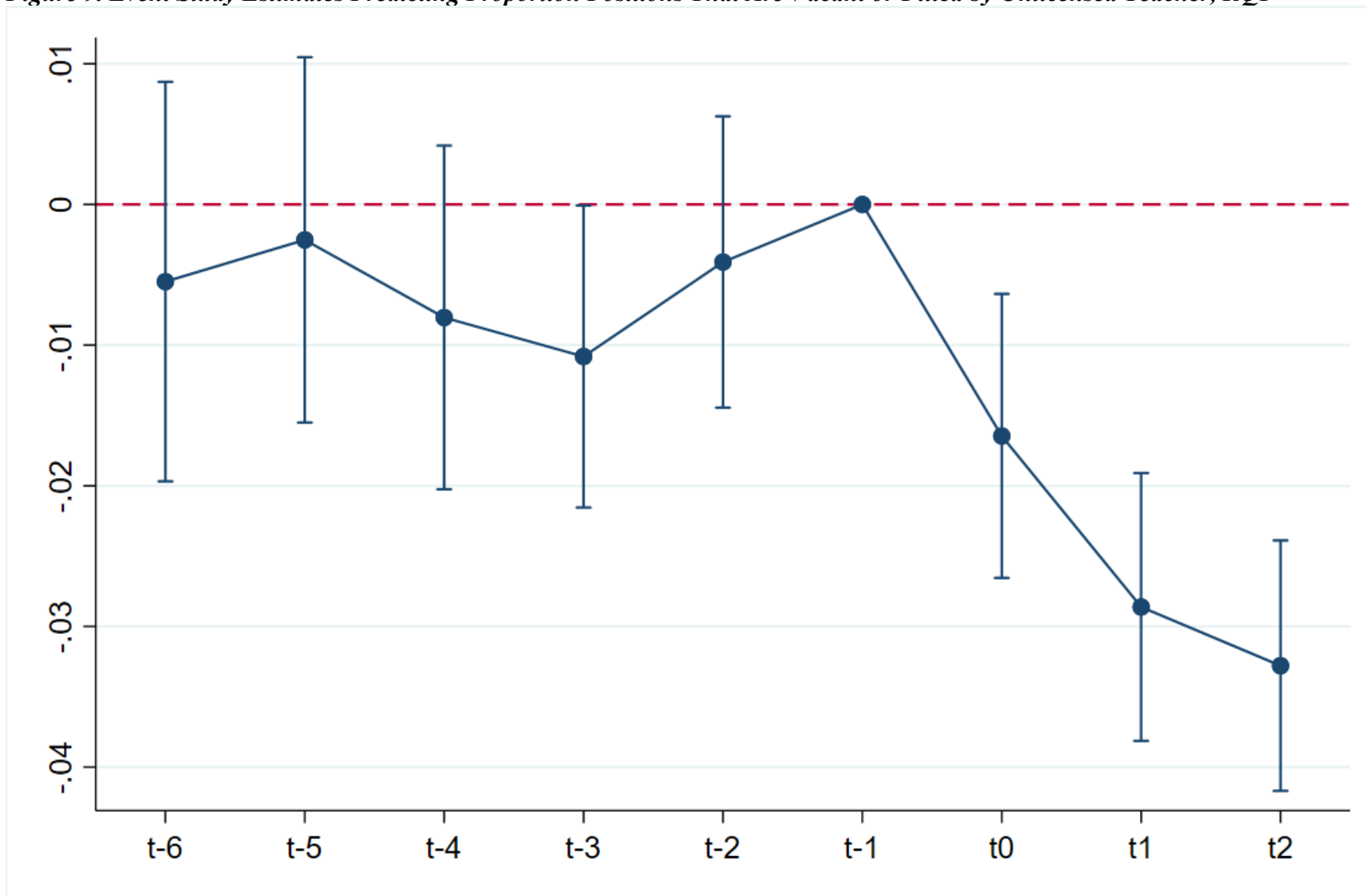
Note. 2015 refers to October 2015 of 2015–16 school year.

*Figure 8. Event Study Estimates Predicting Proportion Positions That Are Vacant, RQ1*



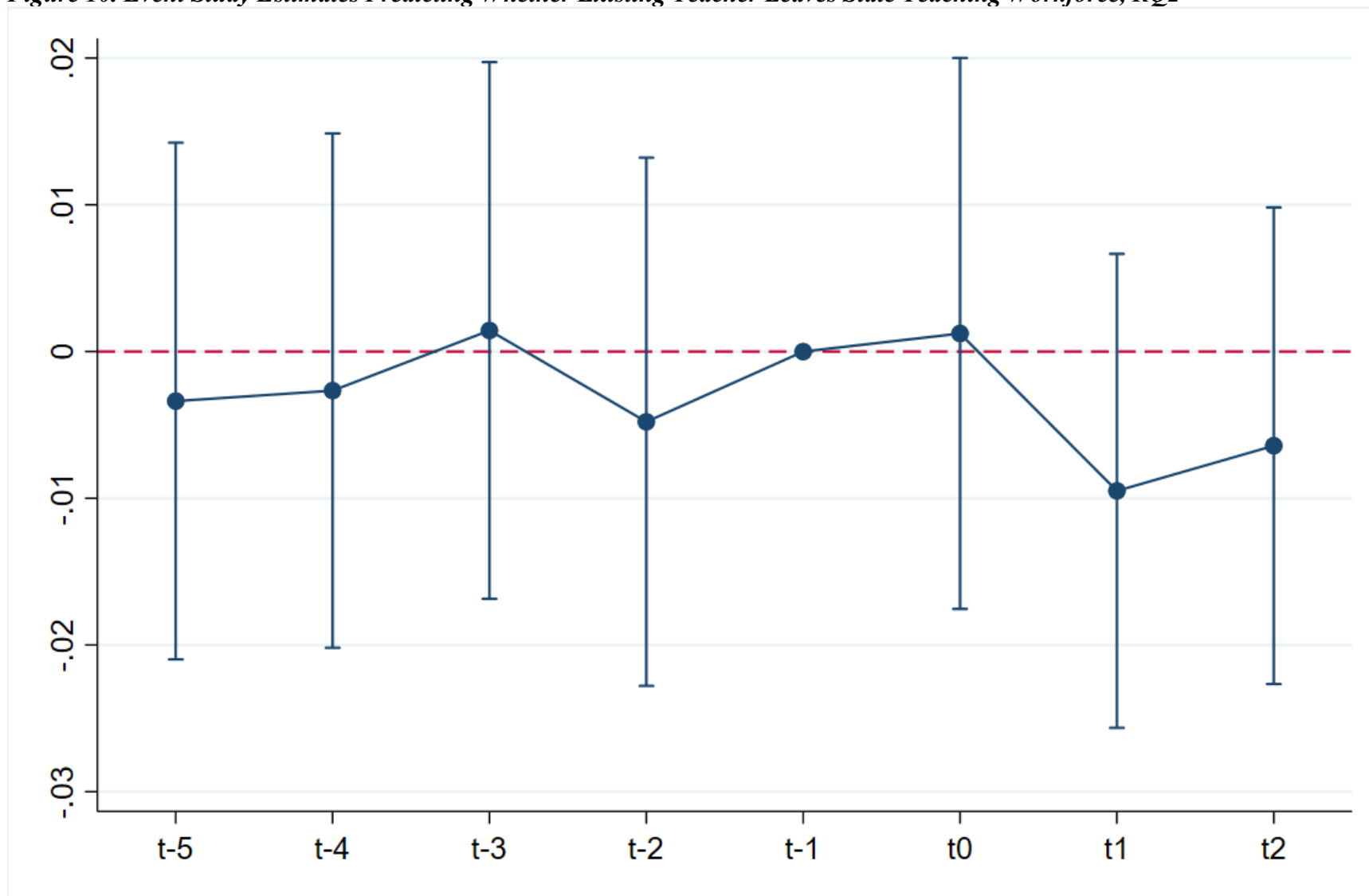
*Note.* Estimates calculated as average marginal effects from event-study logistic regression model.

*Figure 9. Event Study Estimates Predicting Proportion Positions That Are Vacant or Filled by Unlicensed Teacher, RQ1*



*Note.* Estimates calculated as average marginal effects from event-study logistic regression model.

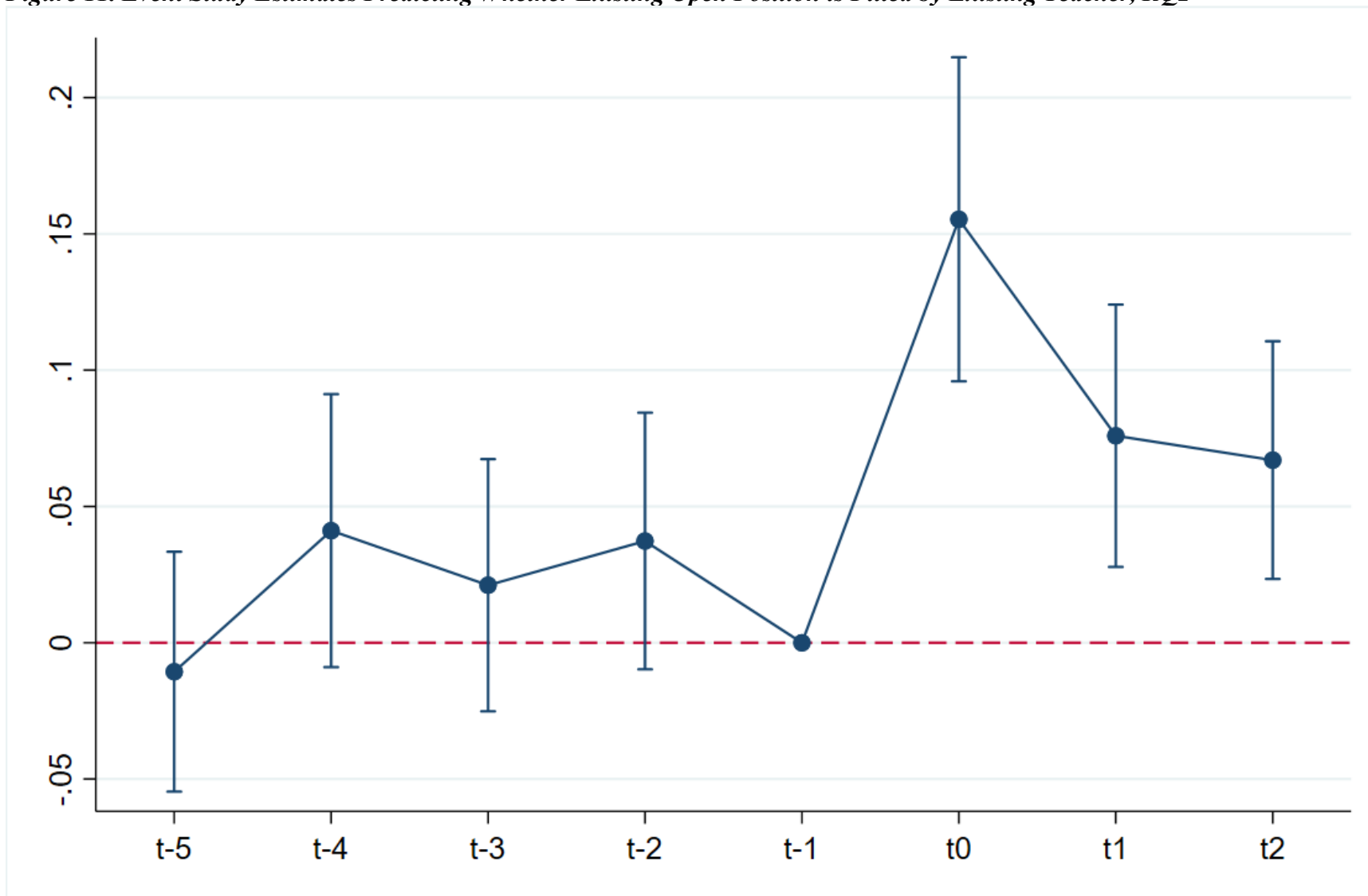
Figure 10. Event Study Estimates Predicting Whether Existing Teacher Leaves State Teaching Workforce, RQ2



Note. Estimates calculated as average marginal effects from event-study logistic regression model.

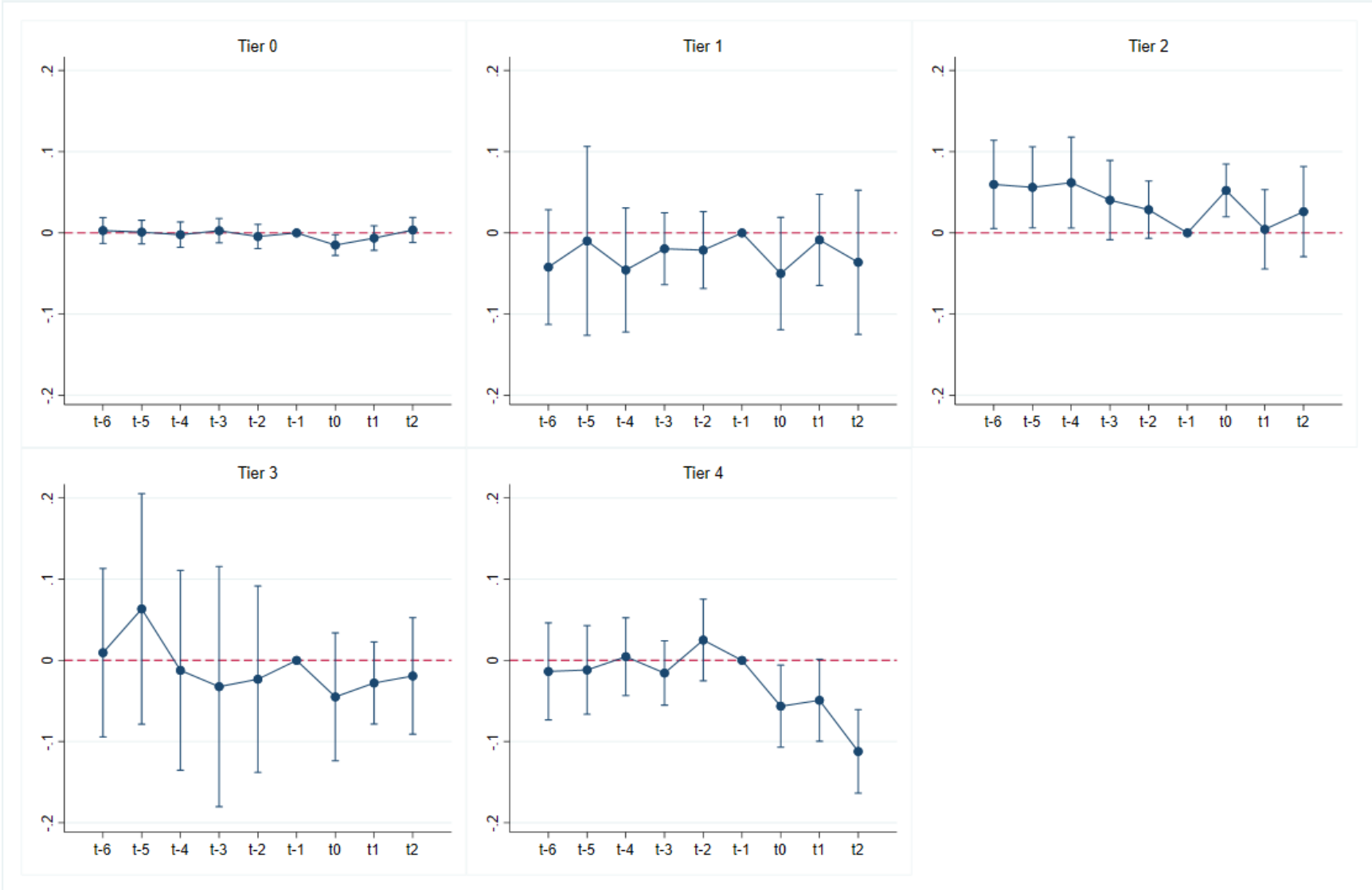


Figure 11. Event Study Estimates Predicting Whether Existing Open Position is Filled by Existing Teacher, RQ2



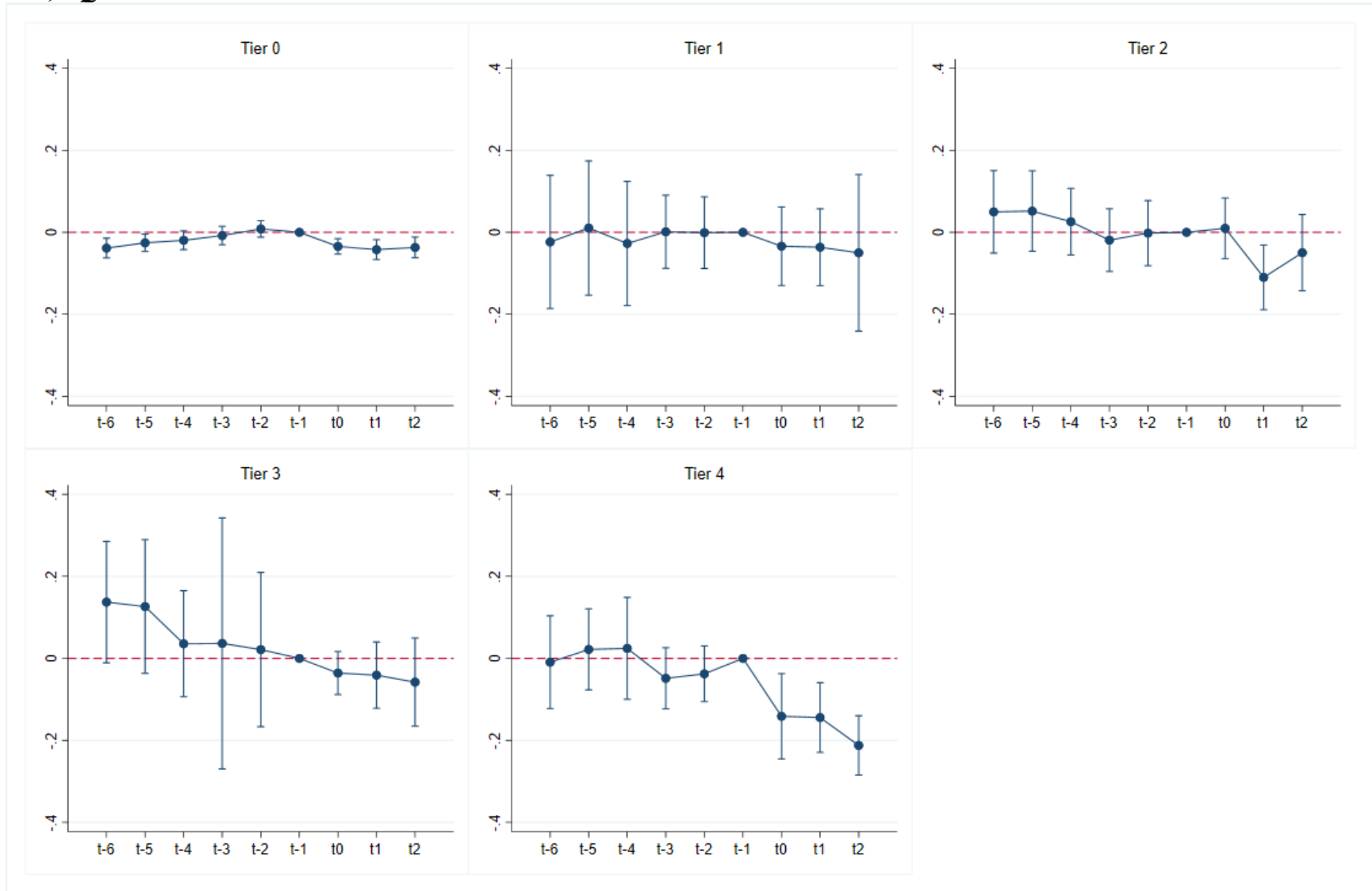
Note. Estimates calculated as average marginal effects from event-study logistic regression model.

**Figure 12. Event Study Estimates Predicting Proportion Positions That Are Vacant by School Tier, RQ3**



*Note.* Estimates calculated from event-study OLS regression models estimated separately by school tier.

**Figure 13. Event Study Estimates Predicting Proportion Positions That Are Vacant or Filled by Unlicensed Teacher, by School Tier, RQ3**



*Note.* Estimates calculated from event-study OLS regression models estimated separately by school tier.

**Table 1. Difference-in-Differences Estimates, RQ1**

	(1)	(2)	(3)	(4)	(5)	(6)
Specification	Logit			OLS		
Scale	Odds ratio	Odds ratio	Marginal effect	Marginal effect	Marginal effect	Marginal effect
<b>Panel A: Outcome = Position is Vacant</b>						
Special Education * Post Bonus	0.675*** (0.069)	0.680*** (0.070)	-0.012** (0.004)	-0.012** (0.004)	-0.013** (0.004)	-0.012** (0.004)
Special Education	2.432*** (0.156)	2.473*** (0.150)	0.027*** (0.002)	0.031*** (0.003)	0.041*** (0.002)	0.041*** (0.002)
Post Bonus	1.309*** (0.080)	1.310*** (0.081)	0.005** (0.002)	0.007*** (0.002)	0.007*** (0.002)	
School FE		X	X	X	X	
School-by-role FE					X	X
School-by-year FE						X
Observations	115,085	113,326	113,326	115,085	115,085	115,085
<b>Panel B: Outcome = Position is Vacant or Filled by Unlicensed Teacher</b>						
Special Education * Post Bonus	0.635*** (0.047)	0.645*** (0.049)	-0.040*** (0.007)	-0.040*** (0.008)	-0.041*** (0.008)	-0.040*** (0.008)
Special Education	2.993*** (0.168)	3.114*** (0.171)	0.079*** (0.005)	0.092*** (0.006)	0.071*** (0.003)	0.070*** (0.003)
Post Bonus	1.163*** (0.044)	1.149*** (0.045)	0.001 (0.002)	0.007*** (0.002)	0.007*** (0.002)	
School FE		X	X	X	X	
School-by-role FE					X	X
School-by-year FE						X
Observations	115,085	114,175	114,175	115,085	115,085	115,085

*Note.*  $p$ -values from two-sided  $t$ -test. \* $p < .05$ ; \*\* $p < .01$ ; \*\*\* $p < .001$ . FE = fixed effect, OLS = ordinary least squares. Outcome in Panel A is binary indicator for whether position is vacant in October of given school year; outcome in Panel B is binary indicator for whether position is vacant or filled by unlicensed teacher in given year.

**Table 2. Difference-in-Differences Estimates, RQ2**

	(1)	(2)	(3)	(4)	(5)	(6)
Specification	Logit			OLS		
Scale	Odds ratio	Odds ratio	Marginal effect	Marginal effect	Marginal effect	Marginal effect
<b>Panel A: Outcome = Existing Teacher Leaves State Teaching Workforce in Following Year</b>						
Special Education * Post Bonus	0.966 (0.048)	0.963 (0.048)	-0.002 (0.005)	-0.002 (0.005)	-0.003 (0.005)	-0.003 (0.006)
Special Education	1.519*** (0.062)	1.510*** (0.062)	0.039*** (0.004)	0.040*** (0.004)	0.040*** (0.002)	0.041*** (0.002)
Post Bonus	1.064* (0.030)	1.066* (0.030)	0.005* (0.002)	0.005* (0.002)	0.005* (0.002)	
School FE		X	X	X	X	
School-by-role FE					X	X
School-by-year FE						X
Observations	99,400	99,372	99,372	99,400	99,400	99,400
<b>Panel B: Outcome = Open Position Filled by Existing Teacher from Other Position Type</b>						
Special Education * Post Bonus	1.979*** (0.238)	2.088*** (0.274)	0.084*** (0.014)	0.081*** (0.014)	0.078*** (0.015)	0.075*** (0.017)
Special Education	1.473*** (0.119)	1.418*** (0.122)	0.061*** (0.007)	0.031*** (0.008)	0.254*** (0.003)	0.087*** (0.003)
Post Bonus	0.866 (0.070)	0.883 (0.075)	0.015* (0.006)	-0.008 (0.006)	-0.009 (0.006)	
School FE		X	X	X	X	
School-by-role FE					X	X
School-by-year FE						X
Observations	14,546	14,102	14,102	14,546	14,546	14,546

*Note.* *p*-values from two-sided *t*-test. \**p* < .05; \*\**p* < .01; \*\*\**p* < .001. FE = fixed effect, OLS = ordinary least squares. Outcome in Panel A is binary indicator for whether existing teacher leaves teaching workforce the following year; outcome in Panel B is binary indicator for whether an open position, defined as a position that is not filled by an existing teacher from the same position type (general education or special education), is filled by an existing teacher from a different position type (general education or special education).

**Table 3. OLS Triple Difference Estimates, RQ3**

	(1)	(2)	(3)	(4)	(5)	(6)
Outcome	Position is Vacant			Position is Vacant or Filled by Unlicensed Teacher		
Special Education * Post Bonus	-0.006 (0.004)	-0.006 (0.004)	-0.006 (0.004)	-0.024** (0.008)	-0.026*** (0.008)	-0.025** (0.008)
Special Education	0.030*** (0.003)	0.038*** (0.001)	0.039*** (0.001)	0.084*** (0.006)	0.066*** (0.003)	0.065*** (0.003)
Post Bonus	0.007*** (0.002)	0.007*** (0.002)		0.008*** (0.002)	0.008*** (0.002)	
Special Education * Post Bonus * Tier 1	-0.003 (0.016)	-0.003 (0.016)	-0.003 (0.015)	-0.009 (0.020)	-0.008 (0.022)	-0.009 (0.020)
Special Education * Post Bonus * Tier 2	-0.007 (0.014)	-0.005 (0.014)	-0.006 (0.014)	-0.042* (0.021)	-0.035 (0.022)	-0.035 (0.023)
Special Education * Post Bonus * Tier 3	-0.026 (0.037)	-0.034 (0.037)	-0.031 (0.037)	-0.080 (0.042)	-0.092* (0.043)	-0.088* (0.041)
Special Education * Post Bonus * Tier 4	-0.066** (0.021)	-0.064** (0.021)	-0.064** (0.021)	-0.134*** (0.034)	-0.134*** (0.034)	-0.132*** (0.034)
School FE	X	X		X	X	
School-by-role FE		X	X		X	X
School-by-year FE			X			X
Observations	115,085	115,085	115,085	115,085	115,085	115,085

Note. *p*-values from two-sided *t*-test. \**p* < .05; \*\**p* < .01; \*\*\**p* < .001. FE = fixed effect, OLS = ordinary least squares. Outcome in columns 1–3 is binary indicator for whether position is vacant in October of given school year; outcome in columns 4–6 is binary indicator for whether position is vacant or filled by unlicensed teacher in given year.