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TEACHER MOBILITY RESPONSES TO WAGE CHANGES: EVIDENCE FROM A QUASI-NATURAL EXPERIMENT

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Teacher mobility responses to wage changes: Evidence from a quasi-natural experiment

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Abstract

This paper utilizes a Norwegian experiment with exogenous wage changes to study teacher's turnover decisions. Within a completely centralized wage setting system, teachers in schools with a high degree of teacher vacancies in the past got a wage premium of about 10 percent during the period 1993-94 to 2002-03. The empirical strategy exploits that several schools switched status during the empirical period. In a fixed effects framework, the wage premium reduces the probability to quit by 6–7 percentage points and increases recruitment by 4–7 percentage points.

1. Introduction

Teachers are important for student achievement, but the evidence indicates that both teacher relative wages and teachers' relative ability have fallen the last decades, see Corcoran et al. (2004), Lakdawalla (2006), and Bacolod (2007). This paper analyzes how wages influence teachers' mobility decisions by exploiting wage variation within a system of completely centralized wage setting.

The decision to stay or leave teaching at a particular school is perhaps the most studied aspect of the teacher labor market. It is attractive from an econometric point of view because at least voluntary quits must be seen as an outcome merely of decisions of individual teachers. In addition, it is reasonable to believe that schools for which teachers tend to quit are in general unattractive for some reason. Determinants of quit decisions are therefore likely to indicate which factors that influence the general attractiveness of schools.

The evidence on how wages affect teachers' quit decisions is mixed.[1] One particular problem with the existing studies is that school district wage levels may respond to teacher behavior. If wages is set in a compensating way and the empirical model does not sufficiently condition on the relevant amenities, one will estimate a negatively biased effect of the wage on teachers' responses. The estimates will be biased in the other direction if wealthy school districts use the wage to attract high-quality teachers. The only previous paper that use a policy intervention to investigate teacher quit behavior seems to be Clotfelter et al. (2008). They exploit a three-year long bonus program in North Carolina public schools serving low-income or low-performing students, and find that the bonus significantly reduced the turnover rate.

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[1] Gritz and Theobald (1996), Murnane and Olsen (1990), Stinebrickner (2001), Boyd et al. (2003) and Ransom and Sims (2008) find negative effects of wages on the probability that teachers leave teaching, and Mont and Rees (1996) find an effect turnover at the school level. However, Eberts (1987) does not find any significant effect on attrition rates, Podgursky et al. (2004) find little evidence that teachers leave teaching for higher pay, and Scafidi et al. (2007) find small and mainly insignificant effects of teacher wages on teacher transitions. Hanushek et al. (2004) find a negative effect of starting teacher salary on the probability that teachers leave the school district, but the effect disappears in models with fixed school district effects.

The lack of instruments for wages makes it attractive to utilize specific institutional settings and experiments with exogenous wage differences to identify teacher behavior. The present paper exploits the wage setting institution for teachers in the Norwegian public-sector primary and lower secondary schools. During a period of more than 40 years, Norwegian teacher wages were solely decided by central wage bargaining. With one exception, the only sources of variation in the wage across teachers have been teaching experience and the amount of formal education. The exception from this rule is the experiment utilized in this paper. In three out of the 19 counties of the country, teachers in schools with a high degree of teacher vacancies in the past got a wage premium of about 10 percent during the period 1993-94 to 2002-03. Within this institutional setting neither teachers, individual schools, or individual school districts could influence the salary, but teachers are expected to react on wage differences in the same way as under local wage setting. The relevant region has in general recruitment problems and vacant teacher positions, and relies on using non-certified teachers in temporary teaching positions.

I estimate both a quit and recruitment model. The recruitment model is motivated by Krueger (1988) and Holzer et al. (1991) who estimate how the number of job applicants responds to the wage. In the present setting where most schools have some vacant teacher positions, the recruitment process is analyzed by relating the probability that a teacher is recently hired to the wage and other characteristics. The quit and recruitment equations describe the supply of teachers directed toward individual schools, which is estimated directly on school level data by Falch (2010).

Systematic factors explaining why some schools have recruitment problems can be hard to observe. One attractive feature of the present experiment is that several schools were eligible for the wage premium only some years during the empirical period. Then it is possible to analyze variation in teacher turnover and wages in a model with fixed school effects capturing the major systematic factors important for low supply.

The paper is organized as follows. The next sections give a description of institutional features relevant for the analysis and the data. Sections 4 and 5 present the quit and recruitment equations, respectively. The results indicate that a wage increase of about 10

percent reduces the probability to quit by 6–7 percentage points and increases recruitment by 4–7 percentage points, which are effects of 20–35 percent. Some stability analyses are presented in Section 6, while Section 7 uses the estimates to calculate teacher supply elasticities at the school level. Section 8 provides some concluding comments.

2. The quasi–natural experiment

To identify teacher behavior, the appointment rules of teachers are crucial as discussed in Bonesrønning et al. (2005). First, In Norway the teachers are linked to the schools and not to the school districts. They are employed by the school district, but the district cannot move the teachers to another school without a major downsizing of the school or an explicit approval by the particular teacher. Second, if at least one certified teacher is interested in the vacant position, a certified teacher will fill the post. According to the school act, a person that is not certified as a teacher can only be employed if no certified teacher applies for a vacant teacher position, and noncertified teachers can only be hired for up to one school year. The subsequent year, the school has to make the vacant position public again in order to search for certified teachers to apply. According to the national regulations, representatives of the teacher union must be informed prior to every hiring decision. In this way the union is able to closely monitor that the schools act in accordance with the law, which has been one of the cornerstones in the teacher trade union policy. Thus, observed teacher shortages in a particular year, defined as the share of noncertified teachers, reflects the state of the teacher labor market that particular year.

The wage determination of teachers is almost completely centralized with basically a common wage schedule at each school in several European countries, for example France, Germany, Italy and the UK. In Norway, the wage of an individual teacher was solely determined by central wage bargaining up to and including the school year 2000–01.[2] The wage varied

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[2] Some limited local wage flexibility was introduced in 2001, and the wage setting was further decentralized from 2004. Centrally decided funds to be used to wage increases for a limited number of teachers became available for the school districts in 2001. The funds may have been used to increase the wage in the schools with lowest supply. On the other hand, local governments may have pursued a strategic policy in which they did not allocate wage increases to schools that could be eligible for extra pay directly from the central government.

across teachers only with respect to education level and teaching experience, but with one exception. The exception was teachers in schools located in one of the three northernmost counties (out of a total of 19 counties) with particular recruitment problems. The eligible schools within these counties were selected by a criterion based on previous teacher shortages, and the central government paid a wage premium of about 10 percent to the certified teachers at these schools. The school districts had no influence on which schools that were eligible of a higher wage and the selection had no financial implications for them.

The experiment includes several schools during the period 1993-94 to 2002-03. Three different systems to reduce teacher shortages have been in place during the empirical period. In 1993-94 to 1995-96, teachers with at least 50 percent of full post in schools with more than 20 percent teacher shortages the past school year received a wage premium in nominal terms.[3] Because the criterion for a higher wage was previous teacher shortages, it was known well in advance of the school year which schools that would be eligible to pay a higher wage.

A new system was in place in the school years 1996-97 to 1997-98. Now only teachers in schools with more than 30 percent teacher shortages in the past school year were eligible for a higher wage. Thus, fewer schools had a wage premium in this period. Within the last system, continuing to the school year 2002-03, teachers in schools with more than 20 percent teacher shortages on average during the four last school years got a wage premium, and the wage premium was equal across all included schools. Falch (2010) includes a closer description of the system.

Since a wage premium is expected to increase teacher supply, schools with teacher shortages marginally above the criterion for paying higher wage are expected to increase the employment of teachers such that the school is not eligible for the wage premium the next school year. In this case, the teachers at the school in 1993–94 kept their wage premium as long as the system was in place, while new teachers at the school did not get the wage

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[3] For schools with 20–30 percent teacher shortages in the past school year, the rules in 1993-94 to 1995-96 differed across school districts. In schools located in one of the north-east school districts in the relevant region, the teacher wage premium was about 10 percent when the shortages last year exceeded 30 percent but only about 5 percent with shortages in the range 20 – 30 percent. In school districts located south-west in the relevant region, the teacher wage premium was about 10 percent in all eligible school districts.

premium. In such semi–included schools, only quits and not hires are expected to differ from other schools. When a new system starts up, and the school is still not eligible to be included in the system, none of the teachers receive the wage premium.

Schools at which teachers have received a wage premium at least once will in this paper be denoted experimental schools. The experimental schools can be in three different states in a particular year; all teachers receive a wage premium (IN-schools), only the incumbent teachers receive a wage premium (SIN-schools), or none of the teachers receive a wage premium (NIN-schools). Several schools switched status during the empirical period. One system-induced explanation is that the requirement to be an IN-school varies over time. In addition, higher wage is expected to increase the supply, and hence, may change the status for a school from IN to SIN or NIN.

The classification of schools was done by state representatives in the relevant counties. Because the criterion for a higher wage was previous teacher shortages, it has always been known in advance whether a school was eligible for the wage premium. From 1998-99 it has been explicit in the instructions from the central government that the classification of IN schools the next school year should be done before March 1. For new positions made public before this date, the school districts have to pay the wage premium without compensation from the state.

Since the classification of schools is based on lagged information, there is no direct causal effect of current teacher behavior on whether there is a wage premium. However, if a school replaces quitting teachers by noncertified teachers, even though there are certified teachers interested in the positions, the incumbent teachers may get a wage premium the next year. There are several reasons why this is not likely to be a problem in the present case. When I collected the data, the state representatives in the counties reported that they did not believe that the system was being manipulated. The appointment rules presented above are crucial in this regard. In addition, the criterion for the wage premium varies over time and has been decided after the registration of teacher shortages. Both the changes in the system in 1995 and 1997 were decided in December, while the criterion is based on teacher shortages earlier in

the fall. Finally, the last system from 1998 and onwards was harder to manipulate since the criteria was based on four year averages.

Anyway, gaming of the system is not expected to influence the estimates of teacher quits in the present paper. Quits are individual decisions, and quitting teachers will not gain from a wage premium at the schools she leaves.

The wage premium was a fixed amount in nominal terms that changed in 1994 and 1998. Thus, the percentage wage premium varied across teachers and across years. The percentage wage premium was lowest in 1993-94 at 7.5 percent and highest in 1998-99 at 12.0 percent.

This experiment differs in several important ways from the North Carolina experiment exploited by Clotfelter et al. (2008). It lasted for a much longer period (10 vs. 3 years), the wage premium was higher (on average 10 vs. 4 percent), all certified teachers were included (in contrast to only math, science and special education teachers in North Carolina), and the system was well known for the teachers (Clotfelter et al., p. 1355, argue that "the vast majority of teachers … misunderstood the provisions of the bonus program").

3. Data

The experimental schools are compulsory primary and lower secondary public–sector schools (first through tenth grade) and are located in three counties consisting of 90 school districts.[4] In the relevant counties, the population is relatively scattered and relatively few workers reside in another school district than their working site.

Individual teacher data with school identifier are provided by Statistics Norway. I use a sample consisting of teachers who have been working at an experimental school at least one year during the empirical period since the identification of the wage premium is based on within-school variation. Since I will estimate models with individual specific effects, the

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[4] Private schools exist to a very small degree. In 1995, only 0.5 percent of the students in the relevant counties were enrolled at private schools.

sample includes all observations of the teachers that have been working at an experimental school. However, observations in which teachers work less than 50 percent of a full-time equivalent position are excluded because they are not eligible for the wage premium. In the merging of teacher and school data, teacher data are missing for two experimental schools established towards the end of the sample period, and school data are missing for two other schools.

Table 1 shows the number of schools and teachers with a wage premium in the different school years. Because some limited local wage flexibility was introduced for the school year 2001-02, only data up to 2000-01 are included in the analysis. Few schools were eligible of a higher wage during the relatively restrictive system in 1996-97 and 1997-98, while about three times as many schools were included in the preceding and following years. The change in the criteria to be included in the system over time implies that most schools only had a wage premium in a short period. Table 2 shows that out of the 161 experimental schools, teachers in 20 schools received a wage premium only in one year and 104 schools received a wage premium in less than four years. 65 schools were eligible for the wage premium only once, but they typically paid a wage premium to incumbent workers (SIN school) in additionally one or two years.

For the relevant counties, the average share of teachers in the empirical period who get a wage premium is shown in Figure 1. School districts in which teachers have received the wage premium are spread all over the region, although the wage premium is most common in the northernmost county. 77 percent of the 90 school districts include at least one experimental school in the empirical period, while 13 percent have an average share of teachers who got a wage premium above 15 percent.

4. The quit equation

4.1. Model specification

A model of turnover behavior is presented in Appendix A. I estimate the linear probability quit model

$$q_{jt}^i = \alpha_j + \alpha_t * \alpha_d + \beta_1 V_{jt} + \beta_2 C_t^i + \gamma (IN + SIN)_{jt+1} + \varepsilon_{jt}^i \qquad (1)$$

where q_{jt}^i is a dummy variable for whether teacher i at school j leaves the school between the school years t and t+1. α_j, α_t, and α_d are school, time, and school district specific effects, respectively, V and C are school and individual charachteristics, respectively, γ is the parameter of interest, and ε is an i.i.d. error term by assumption. The motivation for the extensive set of fixed effects is that they capture characteristics of the choice set of the teachers. The interaction between time and school districts captures all common yearly elements in the teachers' choice set at the school district level. In addition, in the centralized wage setting regime, time fixed effects capture national wage changes.

Mobility costs are expected to vary across individuals. In addition, the valuation of salary and non-pecuniary aspects may depend on individual characteristics as indicated by the model formulations in Appendix A. The empirical model includes a host of individual characteristics as family status and age, which are reported in Appendix B. In addition, I estimate models with teacher fixed effects. The school characteristic included in the model is the number of students.

I will only consider teachers in permanent positions in the quit model in order to focus on the teachers' choices. I relate the quit decision to whether the school pay a wage premium (IN+SIN school) the next school year, consequently relying on quits at the end of the school years 1992-93 to 1999-2000. Descriptive statistics for the sample are presented in Table 3. The experimental schools are relatively small, but the teacher composition does not differ across school types. The average quit rate is 18.2 percent, which is higher than for the population of teachers in permanent positions in Norwegian primary and lower secondary schools reported by Falch and Strøm (2005) and Falch and Rønning (2007). The relatively high quit rate in the present sample seems to be partly related to the fact that experimental schools at the outset are unpopular among teachers, and partly related to the fact that the sample consists of relatively small schools. In addition, in the present sample the quit rate in observations of non-experimental schools is high by construction. Within experimental schools, the quit rate is slightly lower when they pay a wage premium the next school year.

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Table 4 reports mobility patterns, restricted to observations of experimental schools. The first row reflects the quit rates in Table 3. The share of teachers leaving teaching is similar at schools with and without wage premium next year. However, teachers at schools without wage premium next year move to a larger extent to other schools without wage premium.

Figure 2 presents the density of the change in the average quit rate at the school level when a wage premium is introduced and removed. The figure thus takes school specific factors into account. The distribution of the change in the quit rate when a wage premium is introduced is clearly to the left of the distribution when the wage premium is removed. The mean values are -0.04 and 0.09, respectively, indicating an average effect of the wage premium of 0.065.

4.2. Empirical results

The first column in Table 5 estimates the effect of the wage premium on the probability to quit at school using a model including only time and school specific effects. With this simple model formulation, the probability to quit is reduced by 4.8 percentage points when the wage increases by about 10 percent. The next column includes interactions between time fixed effects and school district fixed effects. Then the effect of the wage premium increases to 5.8 percentage points. Thus, school district specific factors do not seem to be highly correlated with the wage premium, even though the sample includes several schools that never pay a wage premium. The model also includes the number of students at school. The result indicates that the quit rate is lowest when schools are small, but the wage effect is not affected by whether school size is included in the model or not.

The model in column (3) in Table 5 includes a range of observable individual characteristics, which does not affect the wage effect. One reason may be that mobility costs are not important, another that mobility costs are captured by the fixed effects in the model. Appendix Table B1 shows that few of the individual characteristics have a significant effect at five percent level.

Individual fixed effects are included in the model in column (4) in Table 5. This specification control for all time-invariant characteristics at the individual level and is by such less vulnerable to omitted variable biases. On the other hand, since the dependent variable is a

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dummy variable, the dependent variable in this within-individual framework only varies for individuals that de facto move in the sample period. Thus, the weight on mobile individuals is higher than in the other models, and these individuals may be the most responsive to external factors.[5] The effect of the wage premium is larger in this model than in the previous models, which indicates that the previous estimates are not biased upwards by excluded individual characteristics.

The wage response estimated is close to Clotfelter et al. (2008). While the effect in column (3) in Table 5 implies a quit elasticity of about -3.3, the results in Clotfelter et al. implies a quit elasticity of about -4.3.

The literature on labor supply finds that women's working hours are more responsive to the wage than men's working hours, see Killingsworth and Heckman (1986) and Pencavel (1998). In addition, Pencavel finds that the wage elasticity is highest for married women and young women. It is not straightforward to translate these supply elasticities at the market level into mobility responsiveness. One could expect that people who respond strongly in terms of hours also respond strongly in terms of mobility. The usual interpretation, however, is that women are more responsive in terms of hours since an attractive alternative for many is to stay at home. But then they are also less geographically mobile, working in the direction of less mobility responsiveness. At the firm level, Manning (2003) finds similar elasticities for both genders, while Hirsch et al. (2008) find that women's labor supply is less elastic than men's.

Column (5) in Table 5 allows for heterogeneous effects of the wage premium across teachers' gender, marital status, and parenthood. The latter interaction is included since it is expected that people are less geographically mobile when they have school-aged children.[6] The results indicate that women and teachers with children only weakly respond to the wage as expected, while the response is largest for married teachers.

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[5] The sample consists of 1810 teachers, for which 289 is observed only in one year. 59 percent of the rest of the teachers quit at least once in the sample period. Using the latter sample, the wage effect increases to 0.090 in the model specification in column (3) in Table 5.

[6] I have also estimated models including interaction between the wage premium and age below 30 years. This interaction effect was, however, clearly insignificant in both the quit equation (t-value of 0.36) and the recruitment equation below (t-value of 0.27).

Table 6 presents the marginal effects of the wage premium for different groups. The marginal effects are computed from models that are more flexible than the model in column (5) in Table 5 in the sense that they allow for different wage responses for married women and married men, married people with and without children, etc. The table shows that three of the eight groups specified respond strongly and statistically significant at five percent level on the wage. That is all groups of married teachers, except married women with school-aged children.[7] For unmarried men, the wage effect is similar to the average effect, but insignificant at five percent level. Overall, there is a distinct difference between female and male teachers. In particular among female teachers it seems like the wage effect depends on family status.

5. The recruitment equation

5.1. Model specification

Quit and recruitment decisions are in general not separable. However, analyzing the matching process of teachers and schools is complicated by the fact that one only observes actual matches and not all potential matches.[8] The present paper seems to be the first to estimate recruitment in the same way as the traditional way to estimate quit decisions. That is possible in the present case since the schools paying a wage premium typically have recruitment problems, markedly reducing the potential influence of other actors (schools, school districts, etc.) than teachers.

I relate the probability that a teacher is recently hired to the wage and other characteristics, including several fixed effects. The advantage of an individual level approach compared to a

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[7] These three groups include 39 percent of the sample. The largest group in the classification of Table 7 is married women with school-aged children, including 20 percent of the sample.

[8] To estimate a structural model including both quit and matching decisions, one needs information on the searching behavior of teachers. Boyd et al. (2003) seems to be the only paper that estimates a matching model for the teacher labor market. They model the initial sorting of newly educated teachers to different schools. By focusing on the initial matching they reduce the number of potential matches and avoid modeling quit decisions. In the present setting the whole labor market for teachers is relevant. Then each individual with a teacher certificate, employed at a school or not, is relevant for vacant teacher posts that exist in the national labor market.

school level approach is the possibility to more properly control for individual characteristics. The recruitment model estimated is

$$r_{jt}^i = \alpha_j' + \alpha_t' * \alpha_d' + \beta_1' V_{jt} + \beta_2' C_t^i + \gamma' IN_{jt} + \varepsilon_{jt}'^i, \qquad (2)$$

with standard errors clustered at the school level. The equation uses the same notation as in equation (1), and r_{jt}^i is a dummy variable for whether teacher i at school j is recruited in year t.

Truncation is a concern in the recruitment model since some schools do not have excess demand. Some schools do not need to recruit noncertified teachers. In this case one cannot rule out that there are teachers interested in working at the school without getting a job offer, i.e., recruitment is not only determined by individual teacher behavior. If wages have a strong effect on teachers' mobility decision, we would expect the wage premium to be positively related to absence of excess demand. In this case, models that do not take truncation into account will underestimate the wage effect on the number of applicants. On the other hand, if excess demand and the wage premium are conditionally positively correlated, which is likely in the present case, the truncation bias will be towards finding no effect.

In the present data, absence of excess demand is the case in only 21 percent of the observations of experimental schools, and does not seem to be correlated with the wage premium. In the sample used in the analysis below, the correlation coefficient between wage premium and absence of excess demand is -0.02. The correlation coefficient for the within-school variation is 0.03. Nevertheless, the relevance of the truncation issue will be investigated in two ways. First, I will estimate models conditional on the share of vacant positions in year t–1, defined as the share of noncertified teachers. With many vacant positions last year, several new teachers need to be hired before excess demand is eliminated. Secondly, I will estimate models restricting the sample to schools with vacant teacher positions. In addition, I will allow the effect of the wage premium to depend on whether the teacher had a permanent teacher position in year t–1. Such teachers only move due to individual preferences.

The probability that a teacher is recently hired is expected to be related to whether the school pays a wage premium the particular year. Consequently, the sample period is 1993-94 to

2000-01. Descriptive statistics is presented in Table 7. The sample used is larger than for the analysis of quits because it includes all certified teachers, not only those with permanent positions. This is also why the recruitment rate is higher than the quit rate. Table 7 clearly shows that the recruitment rate is higher when the experimental schools pay a wage premium to newly hired teachers (IN schools) compared to no wage premium (SIN + NIN schools).

Table 8 shows where the new recruits come from. Schools with wage premium have more teachers coming from non-teaching activity, which includes novice teachers in their first job. They also to a larger extent than schools without wage premium recruit from other schools, including schools in other regions and schools with wage premium. Relatively few of the newly hired teachers come from schools in the same school district.

Figure 3 presents the density of the change in the average recruitment rate when a wage premium is introduced and removed. The distribution when a wage premium is introduced is to the right of the distribution when it is removed, as expected. The mean values are 0.03 and -0.09, respectively, indicating an average effect of the wage premium of 0.06.

5.2. Empirical results

The recruitment equation is more sensitive to model specification than the quit equation. The first column in Table 9 shows that the within school and year correlation is relatively large. When fixed school district effects are included (column 2), the effect of the wage is somewhat smaller. In this model formulation the wage premium increases the probability of being recently hired by 7.5 percentage points, which is significant at one percent level.

Including individual characteristics reduces the effect to 5.4 percentage points. Thus, in contrast to the quit model, individual characteristics are correlated with the wage premium. Table B1 shows that the dummy variable for permanent teacher position last school year has, as expected, a strong negative effect on the probability of being hired at a new school the present school year.[9] When excluding this variable from the model, the effect of interest

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[9] The dummy variable for permanent position last year can also be included in the quit equation if the sample period is restricted to start in 1993-94. In this case, however, the effect of the dummy variable is insignificant at five percent level as expected because the sample in the quit equation only consists of teachers presently in permanent positions.

increases to 7.0 percentage points. Column (4) in Table 9 includes individual fixed effects. In contrast to the quit equation, inclusion of individual fixed effect reduces the effect of the wage premium.[10]

The last column in Table 9 includes the same interaction as for the quit equation. In contrast to the quit equation, the wage response seems to be lowest for married people. Table 10 presents the marginal effects of the wage premium for the different groups. Even though the average wage effect is similar in the quit and recruitment equation, the groups that respond to the wage are different. The wage effect on recruitment is significant at five percent level only for unmarried teachers without school-aged children.[11] For married teachers, the effect is positive but low, and independent of gender and children.

The models in Table 11 are estimated to give an indication of the importance of truncation using the model in column (3) in Table 9 as the starting point. Some schools may have excess supply of teachers. Column (1) in Table 11 conditions on the share of vacant teacher positions last year, which reduces the effect of the wage premium from 5.4 to 4.2. Column (3) restricts the sample to schools with excess demand the present school year and column (5) restricts the sample to schools with excess demand both the present and previous school years. The wage effects in these models are 0.052 and 0.043, respectively. Overall, the results indicate that there is a small positive truncation bias, but the differences are statistically insignificant.

Table 11 also reports results for models where the wage effect is allowed to differ for teachers with a permanent teacher position the previous year. These teachers can stay in their previous position if they want. If they are observed as newly hired the present school year, they certainly prefer the present school over the previous school. I find that they have a lower wage response than other teachers, in the range 2.4–3.8 percentage points, which at five percent

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[10] The sample consists of 2208 teachers, for which 305 is observed only in one year. 68 percent of the rest of the teachers are recently hired at least once in the sample period. Teachers with permanent position last year change school in 12 percent of the observations and teachers with temporary position last year change school in 29 percent of the observations. In addition, in 13 percent of the observations the teacher did not have a teacher position last year and are by definition recently hired. Reducing the sample to teachers that are hired in the sample period, the wage effect increases to 0.088 in the model formulation in column (3) in Table 9.

[11] These two groups, male and female unmarried teachers without school-aged children, consist of 25 percent of the sample. The largest group in the classification of Table 10 is unmarried women without school-aged children, consisting of 18 percent of the sample.

level is insignificantly different both from the effect on other teachers and from zero except in column (2).[12]

6. Stability analyses

Even though the wage response depends on individual characteristics, we would expect constancy across time and regions to the extent that the composition of the teachers does not change. In this section I also test the assumption made so far that the wage responses are static in the sense that it is only the present wage that matter for the individual decisions.

Table 12 present results for both the quit and recruitment equations. The first model formulation allows for different wage effects the first and the last year of wage premium. If it was well known in advance which schools that would pay a wage premium, as intended, the wage effect should not be smaller the first year a school was eligibly of the wage premium as the consecutive years. For the last year with wage premium, the effect on recruitment should be lower when teachers have forward looking behavior. For the quit equation in column (1) in Table 12, the wage response is smallest when wage premium is introduced next year (First year wage premium) and largest the last time with wage premium next year (Last year wage premium), but the differences are not significantly different from the average effect. The wage effect seems to be reasonable stable also in the recruitment equation (column 5), although the point estimates indicate weakest effect the last year as expected.

Columns (2) and (6) in Table 12 allow for different wage effects in the first systems in place, up to 1997-98, compared to the last system. Within the last system the criterion for wage premium is based on average teacher shortages during the last four years and not only the past school year as in the previous systems. If the recruitment effect is biased due to gaming of the system, the effect would be upward biased. Gaming would imply low recruitment of certified teachers in cases without wage premium in order to achieve the premium in the future. Thus, with gaming, the estimated effect of the first systems is expected to be largest. The estimated

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[12] When the sample is restricted to teachers in permanent teacher position last year, the estimated wage effect is 0.044–0.064 for the model specifications in Table 11, and significant at one percent level in all cases.

wage effects on both quits and recruitment are, however, smaller in the first systems than in the last system, although the differences are not significant at five percent level. These results indicate that gaming of the system is at most a minor issue for the present analysis.

Three counties were included in the system. Column (3) indicates that the quit response was lowest in the northernmost county (Finnmark), although the difference to the other countries is insignificant. For recruitment, column (7) indicates that the effect was lowest in the southernmost of the three counties (Nordland). Finally, Table 12 includes a test on whether the wage effect depends on school district size with a cutoff on 1000 students in compulsory schooling (about 70 percent of the teachers are employed in school districts with less than 1000 students), which is rejected.

7. A labor supply elasticity interpretation

At the establishment level, changes in individuals' ranking may be seen as a change in labor supply directed towards the establishment.[13] The supply S faced by a school can be decomposed into current incumbent teachers T, new hires R, and teachers who want to work at the school but who are not offered a post M

$$S_{jt} = T_{jt} + R_{jt} + M_{jt} = \left(1 - q_{jt}\right) N_{jt-1} + R_{jt} + M_{jt}, \tag{3}$$

where N is employment and the quit rate q is a flow variable. Krueger (1988) and Holzer et al. (1991) estimate the elasticity of job applicants with respect to the wage, which can be seen as the wage response of (R + M). It follows from (3) that the short run wage elasticity of labor supply can be written

$$\varepsilon_{Sw} = \frac{\partial S_{jt}}{\partial \ln\left(w_{jt}\right)} \frac{1}{S_{jt}} = \left[\left(\frac{\partial R_{jt}}{\partial \ln\left(w_{jt}\right)} + \frac{\partial M_{jt}}{\partial \ln\left(w_{jt}\right)}\right) \frac{1}{N_{jt-1}} - \frac{\partial q_{jt}}{\partial \ln\left(w_{jt}\right)}\right] \frac{N_{jt-1}}{S_{jt}}. \tag{4}$$

In estimating the recruitment equation above, I assumed that employment is supply constrained, i.e., $M_{jt} = 0$.

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[13] There exist few empirical analyses of overall labor supply faced by individual schools, but Currie (1991), Bonesrønning et al. (2005), and Falch (2010) are exceptions.

A more convenient expression of the supply elasticity can be derived for the steady state where $N_{jt} = N_{jt-1} = S_{jt}$ and $M_{jt} = 0$. Total separations $q_{jt}N_{jt-1}$ is equal to the number of recruits R_{jt}. The short-run elasticity is then given by

$$\varepsilon_{Sw} = \frac{\partial R}{\partial \ln w} \frac{1}{N} - \frac{\partial q}{\partial \ln w} = \left(\varepsilon_{Rw} - \varepsilon_{qw} \right) q_{jt},$$ (5)

while in the long run, setting $S_{jt} = N_{jt-1}$ in equation (3), the elasticity is

$$\varepsilon_{Sw}^L = \left(\varepsilon_{Rw} - \varepsilon_{qw} \right)$$ (6)

One approach in the literature is to assume that separations to and recruitment from non-employment are insensitive to the wage (Manning, 2003, Ransom and Sims, 2008). In this case $\varepsilon_{Rw} = -\varepsilon_{qw}$ in steady state, and the labor supply elasticity can be calculated based on estimates of ε_{qw}.[14] Based on the results in the present paper it is possible to empirically consider the appropriateness of this assumption.

The share of recruits $r = R/N$ is the theoretical counterpart to the empirical specification in this paper. Given constant employment, which is reasonable in the present context since the wage paid by the school district is not affected by existence of wage premium as explained above, the semi-elasticity for the recruitment probability $\frac{\partial r}{\partial \ln(w)} = \frac{\partial R}{\partial \ln(w)} \frac{1}{N}$. Thus, according the assumptions above, $\frac{\partial r}{\partial \ln(w)} = -\frac{\partial q}{\partial \ln(w)}$. The estimated wage effects in Tables 5, 9, and 11 are in the range -0.058 – -0.070 for quits and 0.035–0.075 for recruitment. For the main specifications in columns (3), the effects are -0.058 and 0.052–0.054, respectively, clearly not significantly different. Thus, according to these results, the assumption in the literature that $\varepsilon_{Rw} = -\varepsilon_{qw}$ seems reasonable.

To calculate the supply elasticity, the size of the wage premium needs to be specified. The average wage premium is 9.5 and 9.1 percent in the sample of the quit and recruitment equation, respectively. In the following calculations I assume a wage premium of 0.09 log points.

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[14] Manning (2003) shows that if one relax the assumption that separations to and recruitment from non-employment are insensitive to the wage, the labor supply elasticity can still be calculated without knowledge of the recruitment elasticity by estimating the probability that a recruit comes from non-employment. He finds labor

Using the same estimated wage effects as above and equation (5),[15] the short run labor supply elasticity is equal to 1.0 for the lower estimates and 1.6 for the higher estimates, with 1¼ as the "best" estimate. This result is remarkable similar to the findings in Falch (2010). Falch (2010) analyzes the same experiment for the period 1995-96 to– 2000-01, using school level data on employment, and find a supply elasticity of about 1.4.

It follows from equations (5) and (6) that the difference between short and long run labor supply elasticities depends on the separation or recruitment rates, which are equal in steady state. The quit or recruitment rates are, however, different in the two samples used to estimate the wage effects in the present paper. This is one reason why one should be cautious in calculating long run elasticities in the present case. If one uses average of quit or recruitment rates, the long run supply elasticity is in the range 4.4 – 7.1.

8. Conclusion

Causal evidence on wage effects is hard to establish since in general a myriad of factors influence observed wages. This paper has exploited a system of centralized determined wage differences for teachers in Norway in a fixed effects framework. I find that the effects of a wage premium on separations and recruitment are significant and similar, but not massive. Interpreted in a labor supply framework, the results imply a labor supply elasticity of about 1¼ in the short run.

The empirical results are in accordance with simple economic theory, which predicts that in steady state the separation and recruitment elasticities with respect to the wage are equal in absolute terms. However, I find that the wage effects depend on individual characteristics.

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supply elasticity of about unity. Using the same methodology on German data, Hirsch et al. (2008) find labor supply elasticity of 2–4.

[15] Using equation (5) instead of equation (4) is not exactly right. There are two reasons why the steady state conditions are not exactly fulfilled. First, excess demand is not observed ($S_{it} \geq N_{it-1}$) in 21 percent of the observations, working in the direction of lower supply elasticity. Second, the number teachers increased slightly in the empirical period since the number of students increased, also working in the direction of lower supply elasticity. At the national level, the cohort size increased by about 1.5 percent yearly in the empirical period.

Men respond stronger on wage differentials than women, in contrast to the usual finding for working hours. For marital status, the effects differ between the decision to leave a school or not and the decision to start to work at a specific school or not. Teachers that are married have large quit elasticity but low recruitment elasticity.

The estimated elasticities in the present paper are partial since schools and school districts cannot influence the wage level. The identification is based on centrally determined wage differentials. The general equilibrium effects are smaller in the case of wage spillovers as found by Staiger et al. (2010). When frictions are present in the labor market, as this study suggests, some wage-setting power lies in the hands of the establishments. The possible exploration of monopsony power in the short run must be balanced against long-run considerations when workers can more fully react to wage differentials.

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Table 1. The number of schools and teachers with wage premium

		1993-94	1994-95	1995-96	1996-97	1997-98	1998-99	1999-2000	2000-01
IN	Schools	70	46	34	22	16	63	76	75
	Teachers	405	223	163	66	51	297	362	325
SIN	Schools	0	42	63	0	16	0	15	31
	Teachers	0	266	418	0	59	0	92	218
IN+SIN	Schools	70	88	97	22	32	63	91	106
	Teachers	405	489	581	66	110	297	454	543

Note: IN denotes that all certified teachers at the school receive a wage premium and SIN denotes that only the incumbent certified teachers at the school receive a wage premium.

Table 2. Years of different states for the experimental schools

	0 years	1 year	2 years	3 years	4 years	5 years	6 years	7 years	8 years	Sum
IN	0	65	35	24	13	12	5	3	4	161
SIN	59	44	53	3	2	0	0	0	0	161
NIN	18	10	14	15	12	52	24	16	0	161
IN + SIN	0	20	30	54	14	16	10	5	12	161

Note: IN denotes that all certified teachers at the school receive a wage premium, SIN denotes that only the incumbent teachers at the school receive a wage premium, and NIN denotes that none of the teachers receive a wage premium.

Table 3. Descriptive statistics, quit equation. Mean values (standard deviation)

	Whole sample	Experimental schools	
		Without wage premium next year for incumbent teachers (NIN schools)	Wage premium next year for incumbent teachers (IN + SIN schools)
Quit rate	0.182	0.172	0.157
Number of students	102.8 (103.2)	65.1 (49.5)	64.4 (60.5)
Age	40.9 (10.1)	41.8 (10.2)	41.3 (10.3)
Women	0.601	0.609	0.602
Children 6–18 years of age	0.439	0.453	0.432
Married	0.583	0.616	0.581
Observations	7,867	3,375	2,306
Number of individuals	1,810	1,201	1,059
Number of schools	720	141	158

Note. The sample includes all teachers who have been working in an experimental school during the empirical period (1992-93 to 2000-01), but restricted to observations of permanent position and working at least 50 percent of full time. Sample period is 1992-93 to 1999-2000.

Table 4. Mobility patterns of teachers in permanent position, percent

	Working at a school with wage premium next year	Working at a school without wage premium next year
Do not quit	84.3	82.8
Quit out of teaching	7.6	7.3
Move to a school without wage premium next year in		
the same school district	1.1	2.3
a school district within the relevant counties	0.8	1.3
a school district outside the relevant counties	1.0	1.7
Move to a school with wage premium next year in		
the same school district	0.3	0.2
another school district	0.2	0.1
Observations	2,306	3,375

Note. Same sample as for experimental schools in Table 3.

Table 5. The quit equation.

	(1)	(2)	(3)	(4)	(5)
Wage premium next year	-0.048	-0.058	-0.058	-0.070	-0.076
	(0.013)	(0.018)	(0.018)	(0.022)	(0.029)
Wage premium next year interacted with					
Women	-	-	-	-	0.048
					(0.023)
Married	-	-	-	-	-0.059
					(0.023)
Children 6–18 years of age	-	-	-	-	0.050
					(0.024)
Logarithm of the number	-	0.081	0.069	-0.006	0.074
of students at school		(0.031)	(0.030)	(0.038)	(0.030)
Time FE	Yes	Yes	Yes	Yes	Yes
School FE	Yes	Yes	Yes	Yes	Yes
Time FE * School district FE	No	Yes	Yes	Yes	Yes
Individual characteristics	No	No	Yes	Yes	Yes
Individual FE	No	No	No	Yes	No
Standard error of equation	0.3713	0.3674	0.3608	0.3022	0.3605
Observations	7,867	7,867	7,860	7,860	7,860

Note. The sample is defined in the note to Table 3. The models are estimated by the linear probability model, standard errors that are Huber-White adjusted for heteroskedasticity and clustered at the school level are reported in parentheses. The individual characteristics included in models (3)–(5) are dummy variables for each age; gender; dummy variables for marital status (unmarried, married, divorced, and widow/widower); dummy variables for children below 6 years of age, children 6–18 years of age, children above 18 years of age, and no children; interaction terms between the dummy variables for children and marital status; interaction terms between the dummy variables for children and gender; interaction terms between gender and marital status; dummy variable for whether the teacher is on leave; dummy variable for reduced working hours; dummy variables for years of education; dummy variable for leader position; and dummy variable for whether the teacher works in the same region as born. All coefficients for the model in column (3) are reported in Appendix Table B1.

Table 6. Heterogeneous wage effects in the quit equation

	Married		Non-married	
	Children 6–18 years of age		Children 6–18 years of age	
	Yes	No	Yes	No
Women	-0.009	-0.126*	-0.021	-0.005
	(0.027)	(0.028)	(0.041)	(0.034)
Men	-0.094*	-0.125*	-0.057	-0.067
	(0.033)	(0.035)	(0.053)	(0.037)

Note. The reported heterogeneous effects are calculated based on a modification of the model in column (5) in Table 5. The model is extended by interaction terms between Wage premium next year and (i) Women and Married, (ii) Women and Children 6–18 years of age (iii) Married and Children 6–18 years of age, and (iv) Women, Married, and Children 6–18 years of age. * denotes significance at five percent level

Table 7. Descriptive statistics, recruitment equation. Mean values (standard deviation)

	Whole sample	Experimental schools	
		Without wage premium for new teachers (SIN+NIN schools)	Wage premium for new teachers (IN schools)
Recruitment rate	0.269	0.197	0.254
Number of students	112.6 (110.5)	68.2 (53.1)	58.7 (57.9)
Age	39.5 (10.3)	40.8 (10.4)	40.0 (10.7)
Women	0.612	0.615	0.613
Children 6–18 years of age	0.399	0.418	0.398
Married	0.529	0.558	0.522
Observations	10,482	5,088	1,892
Number of individuals	2,208	1,717	1,202
Number of schools	953	151	158

Note. The sample includes all teachers who have been working in an experimental school during the empirical period (1992-93 to 2000-01), but restricted to observations of certified teachers working at least 50 percent of full time. Sample period is 1993-94 to 2000-01.

Table 8. The origin of the teachers, percent

	Working at a school with wage premium	Working at a school without wage premium
Incumbent	74.6	80.3
Did not teach last year	17.8	14.7
Worked at a school without wage premium last year in		
the same school district	2.6	2.2
a school district within the relevant counties	2.6	1.7
a school district outside the relevant counties	1.3	1.0
Worked at a school with wage premium last year in		
the same school district	0.7	0.1
another school district	0.4	0.1
Observations	5,088	1,892

Note. Same sample as for experimental schools in Table 7.

Table 9. The recruitment equation

	(1)	(2)	(3)	(4)	(5)
Wage premium	0.087	0.075	0.054	0.034	0.079
	(0.016)	(0.016)	(0.015)	(0.018)	(0.022)
Wage premium interacted with					
Women	-	-	-	-	0.007
					(0.019)
Married	-	-	-	-	-0.039
					(0.022)
Children 6–18 years of age	-	-	-	-	-0.022
					(0.021)
Logarithm of the number	-	0.088	0.029	-0.003	0.029
of students at school		(0.034)	(0.028)	(0.036)	(0.028)
Time FE	Yes	Yes	Yes	Yes	Yes
School FE	Yes	Yes	Yes	Yes	Yes
Time FE * school district FE	No	Yes	Yes	Yes	Yes
Individual characteristics	No	No	Yes	Yes	Yes
Individual fixed effects	No	No	No	Yes	No
Standard error of equation	0.4250	0.4118	0.3464	0.3134	0.3463
Observations	10,482	10,482	10,468	10,468	10,468

Note. The sample is defined in the note to Table 7. The models are estimated by the linear probability model, standard errors that are Huber-White adjusted for heteroskedasticity and clustered at the school level are reported in parentheses. The individual characteristics included in models (3)–(5) are dummy variables for each age; gender; dummy variables for marital status (unmarried, married, divorced, and widow/widower); dummy variables for children below 6 years of age, children 6–18 years of age, children above 18 years of age, and no children; interaction terms between the dummy variables for children and marital status; interaction terms between the dummy variables for children and gender; interaction terms between gender and marital status; dummy variable for whether the teacher is on leave; dummy variable for reduced working hours; dummy variables for years of education; dummy variable for leader position; dummy variable for whether the teacher works in the same region as born; and dummy variable for whether the teacher had a permanent teacher position last year. All coefficients for the model in column (3) are reported in Appendix Table B1.

Table 10. Heterogeneous wage effects in the recruitment equation

	Married		Non-married	
	Children 6–18 years of age		Children 6–18 years of age	
	Yes	No	Yes	No
Women	0.033 (0.026)	0.050 (0.029)	0.035 (0.038)	0.101* (0.029)
Men	0.031 (0.036)	0.033 (0.034)	0.080 (0.066)	0.068* (0.031)

Note. The reported heterogeneous effects are calculated based on a modification of the model in column (5) in Table 9. The model is extended by interaction terms between Wage premium and (i) Women and Married, (ii) Women and Children 6–18 years of age (iii) Married and Children 6–18 years of age, and (iv) Women, Married, and Children 6–18 years of age. * denotes significance at five percent level

Table 11. The recruitment equation, sensitive analyses

Sample	(1)	(2)	(3)	(4)	(5)	(6)
	All		Teacher vacancies present year		Teacher vacancies present and last year	
Wage premium	0.042 (0.016)	0.088 (0.028)	0.052 (0.020)	0.090 (0.034)	0.043 (0.022)	0.075 (0.038)
Wage premium interacted with						
Permanent teacher position last year	-	-0.064 (0.026)	-	-0.052 (0.033)	-	-0.044 (0.034)
Logarithm of the number of students at school	0.025 (0.028)	0.025 (0.027)	-0.003 (0.042)	-0.005 (0.041)	-0.024 (0.045)	-0.022 (0.044)
Share of vacant teacher positions last year	0.125 (0.048)	0.123 (0.048)	-	-	-	-
Time FE	Yes	Yes	Yes	Yes	Yes	Yes
School FE	Yes	Yes	Yes	Yes	Yes	Yes
Time FE * school district FE	Yes	Yes	Yes	Yes	Yes	Yes
Individual characteristics	Yes	Yes	Yes	Yes	Yes	Yes
Individual fixed effects	No	No	No	No	No	No
Standard error of equation	0,3463	0.3461	0.3360	0.3359	0.3328	0.3328
Observations	10,468	10,468	7,829	7,829	6,687	6,687

Note. The sample is defined in the note to Table 7. The models are estimated by the linear probability model, standard errors that are Huber-White adjusted for heteroskedasticity and clustered at the school level are reported in parentheses. The individual characteristics are dummy variables for each age; gender; indicators for marital status (married, divorced, and widow/widower); interaction terms between gender and marital status; dummy variables for children below age 18, children below age 6, and children above age 18; interaction terms between the dummy variables for children and marital status; interaction terms between the dummy variables for children and gender; dummy variable for whether the teacher is on leave; dummy variable for reduced working hours; dummy variables for years of education; indicator for leader position; indicator for whether the teacher works in the same region as born; and indicator for whether the teacher had a permanent teacher position last year. All coefficients for the model in column (4) are reported in Appendix Table B1.

Table 12. Stability analyses

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Quit equation				Recruitment equation			
Wage premium next year	-0.069 (0.026)	-0.080 (0.031)	-0.074 (0.039)	-0.077 (0.042)	-	-	-	-
Wage premium	-	-	-	-	0.077 (0.024)	0.073 (0.018)	0.088 (0.024)	0.074 (0.029)
First year wage premium	0.033 (0.026)	-	-	-	-0.012 (0.023)	-	-	-
Last year wage premium	-0.034 (0.045)	-	-	-	-0.034 (0.025)	-	-	-
The wage premium variable interacted with								
Before school year 1998-99	-	0.034 (0.035)	-	-	-	-0.047 (0.031)	-	-
County of Nordland	-	-	0.005 (0.049)	-	-	-	-0.067 (0.035)	-
County of Finnmark	-	-	0.047 (0.047)	-	-	-	-0.027 (0.034)	-
Less than 1000 students in the school district	-	-	-	0.025 (0.046)	-	-	-	-0.027 (0.033)
Logarithm of the number of students at school	0.068 (0.031)	0.068 (0.030)	0.070 (0.031)	0.069 (0.030)	0.029 (0.028)	0.031 (0.028)	0.027 (0.028)	0.029 (0.028)
Standard error of equation	0.3608	0.3608	0.3609	0.3609	0.3464	0.3463	0.3463	0.3464
Observations	7,860	7,860	7,860	7,860	10,468	10,468	10,468	10,468

Note. The model specifications are the same as in columns (3) in Table 5 and Table 9 for the quit and recruitment equation, respectively, except as indicated.

Figure 1. The average extent of wage premiums at the teacher level in the different school districts

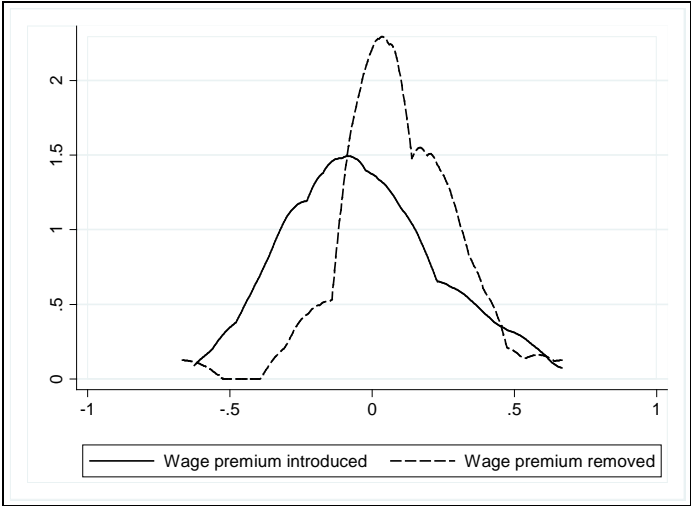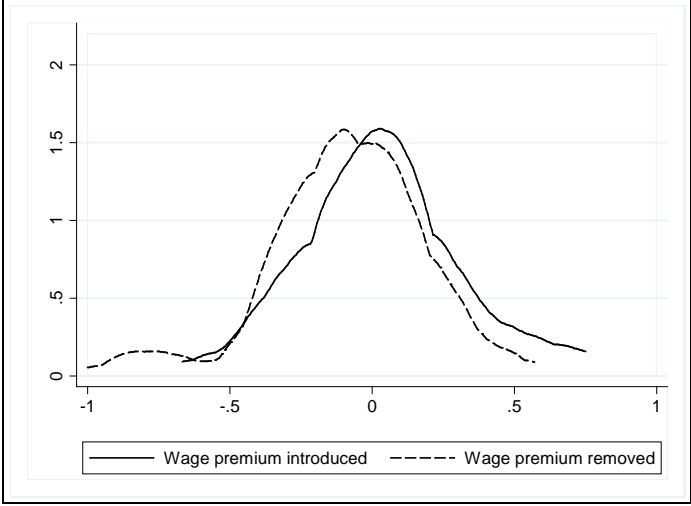Figure 2. Kernel densities for changes in quit ratio at the school level

Figure 3. Kernel densities for changes in recruitment ratio at the school level

Appendix A. A model of teacher turnover behavior

Let the utility of working depends positively on the wage W and other job characteristics V. School characteristics may influence the effort teachers need to provide, and thereby their non-pecuniary rewards. Then the lifetime utility from time t on for teacher i working at j can be written

$$U_{jt}^i = u^i\left(W_{jt}^i, V_{jt}\right) + \delta E\left(U_{jt+1}^i\right), \tag{A.1}$$

where δ is a discount factor, E is the expectational operator, and j is in the choice set J, where J includes schools, other employers, and being out of the labor force. The teacher will prefer another state if

$$u^i\left(W_{jt}^i, V_{jt}\right) + \delta E\left(U_{jt+1}^i\right) < \max_{k \in J}\left\{u^i\left(W_{kt}^i, V_{kt}\right) + \delta E\left(U_{kt+1}^i\right) - c_{jkt}^i\right\}, \tag{A.2}$$

where c_{jkt}^i is the cost of individual i working at j to move to k at time t. Teacher i will prefer to quit j if she is offered a job in k. Assuming $E\left(U_{jt+1}^i\right) = E\left(U_{kt+1}^i\right)$, i.e., the job today does not affect the opportunities in the future in expectational terms, (A.2) can be written

$$u^i\left(W_{jt}^i, V_{jt}\right) < \max_{k \in J}\left[u^i\left(W_{kt}^i, V_{kt}\right) - c_{jkt}^i\right] \tag{A.3}$$

If W_{jt} or V_{jt} increases in a school j, there will be fewer alternatives that yield a higher utility level than staying at the school, which decreases the probability of being offered a job preferable to staying at school j. For teachers not working at school j, the probability that school j is preferable compared to their present position increases. Within this set-up, each factor that makes a school more attractive will both reduce the quit propensity of incumbent teachers and increase the number of applicants to vacant teacher posts. The probability to quit can be written

$$q_{jt}^i = q^i\left(W_{jt}^i, V_{jt}, W_{k \notin jt}^i, V_{k \notin jt}, c_{jkt}^i\right), \quad \partial q_{jt}^i / \partial W_{jt}^i < 0, \tag{A.4}$$

and the number of new hires can be written

$$R_{jt} = R\left(W_{jt}^i, V_{jt}, W_{k \notin jt}^i, V_{k \notin jt}, c_{kjt}^i\right), \quad \partial R_{jt} / \partial W_{jt}^i > 0 \tag{A.5}$$

The dynamic framework of Burdett (1978) gives similar conclusions. In a model with incomplete information and costly job search, the search intensity depends on the present utility level. Burdett shows that higher wages reduce the quit propensity of workers, both because the probability of being offered a more attractive job declines and because it is optimal to reduce the search intensity. At the same time, a vacant position is more likely to be matched with a worker

preferring this job over the alternatives since fewer will reject a job offer. Thus, a rise in the wage will both decrease the quit propensity of existing workers and increase recruitment by higher match probability.

Appendix B. Full models

Table B1. Full models

	Quit equation		Recruitment equation	
	Model in column (3) Table 5		Model in column (3) Table 7	
	Mean [st. dev.]	Coeff. (st. err.)	Mean [st. dev.]	Coeff. (st. err)
Wage premium next year	0.262	-0.058* (0.018)	-	-
Wage premium	-	-	0.164	0.054* (0.015)
Log(Number of students)	4.16 [1.03]	0.069 (0.030)	4.24 [1.05]	0.029 (0.028)
Women	0.601	0.002 (0.042)	0.612	-0.017 (0.033)
Married at present	0.583	0.002 (0.049)	0.530	-0.026 (0.040)
Divorced	0.096	-0.018 (0.065)	0.095	0.074 (0.059)
Widow/widower	0.013	0.374* (0.186)	0.011	0.129 (0.144)
Unmarried	0.308	Ref.	0.364	Ref.
Women * Married	0.348	-0.048 (0.032)	0.323	-0.001 (0.023)
Women * Divorced	0.060	-0.008 (0.049)	0.062	0.012 (0.036)
Women * Widow/widower	0.010	-0.168 (0.112)	0.009	0.005 (0.087)
Children below 6 years of age	0.257	-0.001 (0.041)	0.269	-0.019 (0.037)
Children 6–18 years of age	0.439	-0.054 (0.040)	0.400	-0.010 (0.030)
Children above 18 years of age	0.375	-0.013 (0.061)	0.335	0.001 (0.044)
No children	0.231	0.0003 (0.053)	0.270	0.053 (0.044)
Women * Children below 6 years of age	0.148	-0.030 (0.030)	0.159	0.011 (0.026)
Women * Children 6–18 years of age	0.273	-0.009 (0.022)	0.253	0.009 (0.019)
Women * Children above 18 years of age	0.230	0.029 (0.033)	0.214	-0.005 (0.027)
Women * No children	0.134	0.009 (0.048)	0.158	-0.026 (0.037)
Married * Children below 6 years of age	0.165	0.056 (0.039)	0.164	0.047 (0.033)
Married * Children 6–18 years of age	0.325	0.036 (0.043)	0.283	0.002 (0.030)
Married * Children above 18 years of age	0.288	-0.0011 (0.055)	0.251	0.026 (0.039)
Married * No children	0.043	-0.051 (0.053)	0.041	-0.016 (0.045)
Divorced * Children below 6 years of age	0.010	0.007 (0.067)	0.011	0.024 (0.051)
Divorced * Children 6–18 years of age	0.045	0.061 (0.058)	0.044	-0.002 (0.048)
Divorced * Children above 18 years of age	0.060	0.007 (0.072)	0.059	-0.042 (0.054)

Divorced * No children	0.008	-0.001 (0.090)	0.009	-0.154 (0.087)
Widow/widower * Children below 6 years of age	0.0004	0.580 (0.350)	0.0003	0.209 (0.216)
Widow/widower * Children 6–18 years of age	0.005	-0.155* (0.078)	0.004	-0.112 (0.085)
Widow/widower * Children above 18 years of age	0.010	-0.219 (0.132)	0.009	-0.033 (0.086)
Widow/widower * No children	0.001	0.265 (0.194)	0.001	0.375 (0.360)
On leave	0.008	-0.055 (0.052)	0.009	-0.070* (0.028)
Reduced working hours	0.133	0.031 (0.016)	0.148	0.011 (0.012)
4 years of higher education and non-leader position	0.373	0.025 (0.013)	0.401	0.020 (0.011)
5 years of higher education and non-leader position	0.173	0.028 (0.017)	0.175	0.022 (0.014)
More than 5 years of higher education and non-leader position	0.012	-0.024 (0.046)	0.015	0.011 (0.029)
Years of higher education missing and non-leader position	0.001	0.305* (0.155)	0.002	0.115 (0.098)
Leader position	0.156	0.004 (0.015)	0.139	-0.001 (0.013)
3 years of higher education and non-leader position	0.285	Ref.	0.268	Ref.
Work in the same region as born	0.353	-0.040* (0.012)	0.342	-0.012 (0.009)
Born region missing	0.081	0.004 (0.020)	0.065	-0.003 (0.015)
Working in another region than born	0.566	Ref.	0.593	Ref.
Permanent position last year	-	-	0.699	-0.441* (0.015)
Dummy variables for age	-	Yes	-	Yes
School fixed effects	-	Yes		Yes
Time – School district interaction fixed effects	-	Yes		Yes
Standard error of equation		0.3608		0.3464
Observations		7,860		10,468

Note. The models are estimated by the linear probability model, standard errors that are Huber-White adjusted for heteroskedasticity and clustered at the school level are reported in parentheses. * denotes significance at 5 percent level.