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Child-to-Teacher Ratio in Day Care and Teacher Sickness Absenteeism

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Preface

This paper is part of a series of three independent papers which look into the role of work pressure for day-care teachers' labour-market situation. In this paper, we focus on day-care teachers' sickness absence. Another paper looks at teacher turnover. And the third paper focuses on voluntary early retirement. In all analyses, work pressures are measured by the child-to-teacher ratio. We have benefited from discussions with Thomas Lund and from comments and suggestions from Jan Høgelund, Jacob Nielsen Arendt and Eskil Heinesen. Finally, we are grateful for valuable inputs from an expert monitoring group from ministries, organisations and research institutions with particular knowledge of work environment and labour-market situation of day-care teachers.

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Mette Gørtz
June 2010

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Summary

This paper analyses whether work pressure measured by the child-to-teacher ratio, i.e. the number of children per teacher in an institution, affects teacher absenteeism due to sickness in Danish day-care institutions. We control for individual teacher characteristics like for instance education and family background. Furthermore, we investigate the role of other characteristics at the workplace level like for instance the size of the institution, the proportion of the staff who is trained preschool teachers, family background characteristics of the children in the preschool etc. Our estimation results indicate that for preschool teachers, the risk of becoming long-term ill is positively related to the child-to-teacher ratio for 2005-2006, but not for 2002-2004. We find no significant relationship between the child-to-teacher ratio and the risk of long-term sickness absence for nursery-care teachers. Moreover, we find no significant relationship between the length of long term absence periods and the child-to-teacher ratio. Furthermore, we look at how the extent of short-term absence is related to the child-to-teacher ratio. Our estimation results indicate that nursery-care teachers' sickness absence is positively related to their work pressure, as suggested by our theoretical model, but for preschool teachers, the relationship is insignificant.

1 Introduction

Sickness absence is an important cost to society. This cost encompasses both direct costs to health care, sickness benefits to ill employees, costs associated with substitute employees etc. and indirect costs related to a reduction in productivity at the firm level. Moreover, long and repeated periods of sickness absence make the return to employment difficult, and many long-term sickness periods enhance the risk of ending up on disability pension. Employees on sick leave may experience an impairment of their human capital development at the workplace, reduced career options and a transient – or permanent – dip in the future wage profile. Sickness absence of an employee may also affect the productivity and labour-market outcomes of colleagues at the workplace as well as the family of the sick-listed, cf. Tompa (2002).

This paper focuses on sickness absence of employees in Danish day-care institutions, i.e. nursery care and preschool. Absence due to sickness in the day-care sector incurs high direct and indirect costs and is a source of detriment to the quality of day-care facilities in Denmark.

The paper combines insights from the literature on health economics with the literature on occupational health. The theoretical starting point of the paper is Grossman's model of health capital (1972; 2000). Based on the Grossman model, the paper sets up an empirical model of sickness absenteeism. The paper aims at identifying the main explanations of individual sickness absence of teachers in Danish day-care institutions. The literature on occupational health points at pressure of work as a trigger of absence due to sickness, cf. Lund et al. (2005). One potential contributor to pressure of work in day-care institutions is the child-to-teacher ratio, since a high number of children per teacher may enhance the work pressure for the personnel. The child-to-teacher ratio varies over time, across municipalities and possibly also across institutions within municipalities. In particular, the paper investigates the role of municipal level of child-to-teacher ratios for the incidence and duration of sickness absence in Danish day-care institutions. The empirical analysis differentiates between explanations at the individual level and correlations at the workplace level. We control for background characteristics of the staff in preschool and nursery care and for characteristics of the workplace/firm/institution and the municipality. A very unique feature of the data is that we can identify the children who are connected to the institutions/firms in our sample. This allows us to investigate whether family background of the children in day-care institutions has any impact on sickness absence of the employees, for instance if social problems in the child group enhance the work pressure of the employees in certain institutions.

The main contribution of the paper is that we have access to objective and comparable information on work pressure measured by the child-to-teacher ratio. We exploit the panel dimension of the data as well as instrumental variables to account for the possible endogeneity of the child-to-teacher ratio.

2 Previous Evidence

Previous research in the occupational health literature points to work pressure as a trigger of absence due to sickness. Based on Danish data, Labriola et al. (2006) find that 40 per cent of absence due to sickness can be attributed to workplace conditions. The work environment may affect sickness absence through a number of different channels, see for instance Benavides et al. (2001), Lund et al. (2005) and Vahtera and Kivimäki (2001). According to this line of literature, a poor working environment can 1) directly cause illness or stress; 2) enhance employees' liability to catch ordinary diseases like for instance a cold; 3) have a demotivating effect. Moreover, a poor working environment can affect mobility out of the firm, the sector or the labour market and lead to for instance early retirement.

Afsa and Givord (2006) empirically investigate the effect of working irregular schedules on sickness absence for male manual workers. They reduce the problem of selectivity bias (selection into irregular working schedules) by using propensity-score matching. They find that working irregular hours has a significant impact on sickness absence, especially for older workers.

Lindeboom and Kerkhofs (2000) observe that sickness absenteeism of Dutch public-school teachers varies considerably across schools. They find strong effects of both observed personal characteristics and school characteristics. Moreover, unobserved workplace-specific effects largely take into account the observed variation of sickness absenteeism across schools. These results confirm the more general finding that peer pressure in teams affects employee behaviour, see Mohnen et al. (2008). Hence, peer effects from colleagues may be important. These may be reflected in "cultural" (unobserved) differences across institutions and preschools in attitudes to and habits of reporting sick. Clotfelter, Ladd & Vigdor (2007) analyse the incidence of teacher absences using detailed data from North Carolina. They find that the incidence of teacher absences is regressive in the sense that schools serving low-income pupils would have a higher incidence than schools serving high-income families. Moreover, absences are associated with lower student achievement in elementary grades.

Johansson & Palme (1996) use a sample of Swedish blue-collar workers to investigate the importance of economic incentives for work absence. They model absenteeism as an individual day-to-day decision where absence depends on, among others, the individual wage rate as well as compensation rate in the case of being absent due to sickness. The model builds on the efficiency wage hypothesis taken from Shapiro and Stiglitz's (1984) efficiency wage ("shirking") model that states that when the possibilities of monitoring worker's job performance are poor, employers may pay wages above the market-clearing level in order to obtain adequate effort from their employees. Johanson & Palme (1996) find that the direct cost of being absent has a negative effect on sickness absence for men.

Ose (2005) also uses the "shirking" model of Shapiro and Stiglitz to develop a model of worker absenteeism with working conditions explicitly included. Ose uses Norwegian firm-level data for the 1990s to analyse the importance of working environment factors measured in a special survey which was filled in by a total sample of 331 firms. Ose finds that physical surroundings in the working area, work strain and "cultural factors" like for instance cooperation among employees, relationship of trust between employees and the closest superior and the potential for employees to influence their own work situation have a significant effect on firms' average sickness absenteeism.

There are large costs, both direct and indirect costs, associated with absenteeism. Pauly et al. (2002) find that the costs of lost work time are high when substitutes to replace absent workers are not readily available. In that case, absenteeism may affect not only the worker and the firm, but also the colleagues working with the worker in a team production. Berger et al. (2001) underline the need for analysing further the complicated relationship between productivity and health, and the complex role for employers in terms of promoting employee health and hence productivity in the firm. Sickness among personnel is costly for firms in terms of replacement and reduced productivity for sick workers and hiring and learning costs for new workers, if employees become permanently disabled or find other jobs. Firms' interest in providing health-promoting facilities to workers is higher if the benefits of these can be internalized in the firm as may be the case with low-risk working conditions, see Currie & Madrian (1999) for a survey.

3 Theoretical Background

We use Grossman's model of health capital (Grossman 1972; 2000) to develop a model of how work pressure affects sickness absence. Grossman's model draws heavily on both human capital theory, cf. Becker (1964), Mincer (1974), and on Becker's theory on the allocation of time and his household production theory, cf. Becker (1993.). Although much of our understanding of health capital is inspired by human capital theory, Grossman (2000) points out that health capital differs from other forms of human capital in the following way: While a person's stock of knowledge affects his market and non-market productivity, his stock of health rather determines the total amount of time he can spend producing money earnings and other commodities. Grossman's model of health capital has been extended and refined in a number of theoretical and empirical applications, see for instance Muurinen (1982), Wagstaff (1986; 1993). According to Grossman's health capital model, an individual's health stock is perceived as a capital good, which evolves over time and is subject to depreciation. Individuals invest in their health stock for two main reasons. First, individuals receive direct utility from their health stock in the form of "healthy hours"; this is referred to as the consumption motive for health. Secondly, the health stock is important for productive reasons, since good health is important for labour-market outcomes (and hence income for buying consumption goods) and for the individual's productivity in home activities; this mechanism is usually referred to as the investment motive. According to Grossman, health serves both as a direct input into the utility function (as a consumption good) and determines income and wealth in a life-cycle context (as an investment good). Thus, individuals derive utility in each time period from both directly consuming health "services" in the form of healthy hours from the stock of health and from consuming a composite commodity, which is produced with inputs of market goods and home time (household production and leisure). The intertemporal utility function of a consumer is formulated as:

$$U = U(h_t, C_t)$$

where $h_t = \upsilon_t H_t$. H_t is the stock of health at age t or in time period t, υ_t is the service flow per unit stock and h_t is total consumption of health services in period t. C_t is consumption of another commodity and is produced with inputs of market goods X_t and home time, T_t:

$$C_t = C(X_t, T_t; E_t)$$

E_t, which is the consumer's stock of knowledge or human capital exclusive of health capital, is assumed to be exogenous or predetermined, as we are generally examining individuals after completion of education. An increase in knowledge capital is assumed to raise efficiency of the production process in the non-market or household sector, in the same way as it is usually found in the market sector. By definition, net investment in the stock of health equals gross investment minus depreciation:

$$H_{t+1} - H_t = I_t - \delta_t H_t$$

where I_t is gross investment and δ_t is the rate of depreciation during the t^{th} period, ($0<\delta_t<1$). In Grossman's original model, the rate of depreciation is exogenous, but depends on age. In our paper, we assume in addition that the rate of depreciation depends on work pressure, π_t. This assumption is in line with Sickles and Taubman (1986).

$$\delta_t = \delta(age_t, \pi_t; E_t)$$

Depreciations in health capital are positively correlated with age. π_t symbolises workplace characteristics including work pressure.[1] An increase in the rate of depreciation due to for instance aging diminishes the optimal level of health capital. Consumers produce gross investment in health using inputs of medical care, M_t, and time inputs, TH_t, which may cover time spent going to the doctor, but also preventive actions like exercising etc.:

$$I_t = I(M_t, TH_t;\ E_t)$$

Individuals maximise the present value of their lifetime utility subject to the present value of their lifetime budget constraint, which includes labour earnings and public transfers (sickness benefits), and a per-period time constraint, where Ω is the total amount of time available in any period:

$$TW_t + TH_t + T_t + TL_t \equiv \Omega$$

where TW_t is market work and TL_t is time lost from market and non-market activities due to illness. Consequently, sick time is inversely related to the stock of health, that is $\partial TL_t / H_t < 0$.

Individuals maximise utility subject to a life-time budget constraint and a per-period time constraint. Most empirical applications choose either the consumption or the investment model in order to derive closed-form expressions of the model. In this paper, we want to accommodate both motives for keeping a certain stock of health. Therefore, based on the general model of health capital outlined above, we derive an inverted "demand" function for hours lost due to sickness. The corollary deducted from the above is that a deterioration in working conditions, for instance an increase in the work pressure, π_t, will enhance the level of depreciation of health capital, reduce the optimal level of health and increase the number of hours sick. Moreover, higher age is associated with a higher level of sickness absenteeism. The direction of the effect of the wage in the model is indeterminate. In the pure investment model, a higher wage reflects a higher marginal product in market production which will motivate individuals to invest more in their health. Thus, number of hours lost due to sickness is lower for individuals with a higher wage, reflecting a high opportunity cost of being ill. In the pure consumption model, the direction of the wage effect is indeterminate and depends on the relative marginal product of time in health production versus market production, respectively. Hence, individuals may choose (more or less voluntarily) to leave the labour market to spend more time on health, if the shadow value of their time spent in the labour market is relatively low.

$$TL_t = f(\delta(age_t, \pi_t;\ E_t), W_t;\ E_t, Z_t)$$

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[1] The positive correlation between age and depreciation in health is crucial in Grossman's model for ensuring finite life and leads to an endogenous specification of length of life. Consequently, a person dies when his health stock falls below a threshold level of health.

Z_t symbolises household characteristics. Education (human capital) affects TL_t both through the human capital effect on productivity in health production and through a lowering of the depreciation rate of health capital. An attractive feature of this model is that TL_t can be determined without measuring the health status.[2]

[2] Currie & Madrian (1999) discuss the "measurement bias" in measuring health, see also Bound et al. (1999).

4 Data

The data set we employ for the empirical analysis is a large micro panel based on administrative registers from Statistics Denmark. The population in the data is defined as all individuals who were employed in Danish preschools and nursery-care institutions sometime over the period 2002-2006. We focus on this period for two reasons. First, the definition of personnel resources in the official statistics of Statistics Denmark changed around 2001. Secondly, the reporting of long-term sickness absence for public-sector employees changed around 2000. Due to a change in reporting of the number of children enrolled in day-care institutions, a structural break in the child-to-teacher ratio reporting occurs between 2003 and 2004. Therefore, the analysis is carried out separately for the periods 2002-2004 and 2005-2006. This is explained further in section 5.

We focus on employees in day nursery ("*vuggestue*") and preschool ("*børnehave*"), see section 5. The registry gives us information on around 20,000 pedagogical employees in nursery and preschool each year, of which around ¼ work in nursery care. Within this sample, around 2/3 of the personnel are trained day-care teachers, while 1/3 work as assistant day-care teachers. We explain more about the Danish day-care sector in section 5.

The data set also contains information on labour market status for this population, i.e. for each year we know sector of employment, institutions/firms, unemployment periods, retirement or other periods outside the labour force, periods of long-term absence (more than two weeks) due to sickness, maternity leave etc. Thus, we can track employees' transitions into and out of the day-care sector, and we can also analyse mobility across institutions/firms within the sector. Apart from information on long-term sickness (i.e. periods of absence longer than two weeks), we have access to information at the individual level on total absence due to sickness (i.e. both long-term and short-term spells) for 2005 and 2006 for around 2/3 of the employees. Moreover, we have information on the employees' use of health care, i.e. visits to general practitioners, specialists, hospitalisation, medical (somatic) diagnoses etc.

A unique feature of the data is the possibility to identify the children who are enrolled in the specific day-care centres. Due to data limitations, it was not possible for Statistics Denmark to connect all day-care institutions to one workplace. Approximately 53% of the day-care teachers are connected with their particular child group. Based on this subset of the data, we investigate whether the family background of the enrolled children has any impact on teacher turnover. This could be the case if for instance social problems among the children enrolled in specific day-care centres enhance the work pressure of the employees.

5 The Danish Day-Care Sector

Danish day-care institutions are run by the municipalities, and the cost of day care is largely borne by municipalities. Parents pay a user fee per child which accounts for 20-30% of the total cost. Municipalities can decide on the level of the parents' share, but the maximum parental share is set by law to around 33%. Day-care institutions are directed at preschool children aged ½-6 years. Day-care institutions constitute of day nursery (child-minding), preschool and age-integrated institutions (institutions with children aged ½-6 years, where day nursery and preschool are combined in one institution). At the national level, there were around 300 nursery care institutions in 2006, almost 1900 preschools and a comparable number of age integrated institutions. Municipalities organise their day-care institutions quite differently, when we look at the allocation of children in different types of day care. In some municipalities, the majority of day-care institutions is in the form of age-integrated institutions (age ½-6 years). In other municipalities, the majority of children aged 3-6 are in traditional preschools. For the group of children aged ½-3 years, some municipalities offer nursery-home openings, whereas other municipalities rely on municipally organised child minding.

In the analysis in this paper, we focus on institutions that are either nursery-care institutions or preschools. We disregard child minding and age-integrated institutions. Age-integrated institutions are excluded since they serve both infants below 3 years of age and preschool children aged 3-6. Thus, variation over municipalities and over time in the child-to-teacher ratio of age-integrated institutions may arise as a result of variation in the age distribution of the children in a given year or across municipalities. Moreover, there are no figures for the child-to-teacher ratio in child minding.

The level of the child-to-teacher ratio is decided by the municipality, but there may be some variation within the municipality. This analysis uses municipality averages of the child-to-teacher ratio. The level of child-to-teacher ratio is measured as number of children per full-time day-care employee who are occupied with child care (teachers and assisting teachers with pedagogical functions). Thus, staff occupied with kitchen duties, cleaning, maintenance and repair etc. are not part of the child-to-teacher ratio. Both the number of children and the staff is registered at a certain date. Information about the number of children is from Statistics Denmark's survey on day-care institutions. Up until 2003, the number of children in day-care institutions was registered in the spring. From 2004 and onwards, the children were counted in the autumn. This change in the date of the day-care survey gives rise to a discontinuity in our child-to-teacher ratio measurement. The number of (full-time) employees in day care is based on a Statistics Denmark survey of employees counted in November. This figure includes substitute teachers if the regular teacher is on maternity leave. Thus, there is a risk of double counting leading to inflated staff figures. On the other hand, the survey does not reflect vacancies.

On average, the child-to-teacher ratio in nursery-care institutions was 3-3.5 children per teacher or assistant teacher over the period 2001-2006, see figure 5.1. In preschool (i.e. institutions for children aged 3-6 years), the child-to-teacher ratio was 6-7 children per pedagogical employee over the period 2001-2006, see figure 5.2. There seems to be a tendency that municipalities have a fairly constant child-to-teacher ratio over time.

Figure 5.1 Child-to-teacher ratio in nurseries, 2001-2006

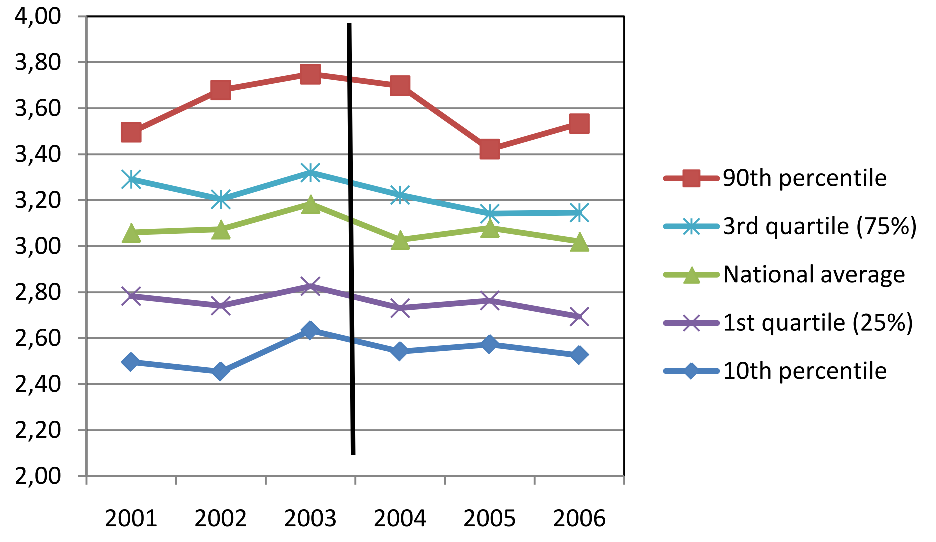
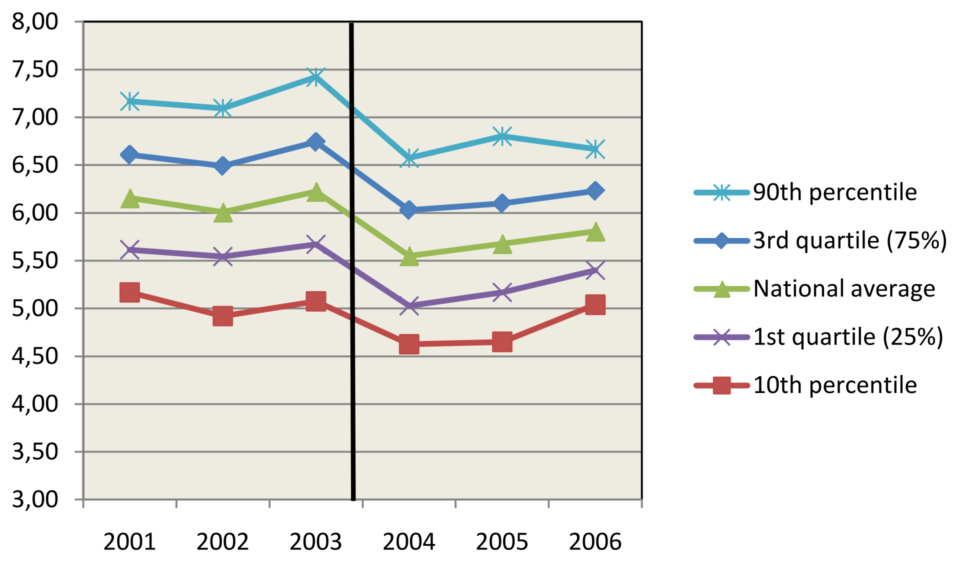


Figure 5.2 Child-to-teacher ratio in preschool, 2001-2006



Note: Figures 5.1 & 5.2: For 2001-2003, the number of both employees and number of children enrolled in day-care institutions was reported in March, while for 2004-2006, the reporting was carried out in September. Thus, the child-to-teacher ratio cannot necessarily be compared across these two periods.

Consequently, the empirical analysis focuses on day nursery (children aged ½ up till 3 years) and preschool (children aged 3-6 years). Day-care teachers have completed a 3½-year pedagogical education. Assisting day-care teachers have usually followed a 2-year education with a combination of school and practice teaching. Day-care institutions also employ staff with no education. These are typically young people who gain labour-market experience before starting further education or older employees with on-the-job practice in the sector. On average, trained teachers account for almost 50% of the employees in day-care institutions, while assistant teachers amount to almost 40% of all employees.

The work pressure for preschool teachers may not only depend on the number of children, but also on the pedagogical challenges in the child group. Thus, it is possible that children's parental background has implications for the work pressure in the day-care institution. On the one hand, a high proportion of disadvantaged children might induce more sickness among staff (directly or through selection). On the other hand, a high proportion of disadvantaged children might promote dedication to work and might attract teachers with a high level of devotion to and engagement in their work. Both of these explanations might be at work simultaneously, and the direction of the net effect is mainly an empirical question. Our prior is that family background is important for children's social abilities and thereby their educational attainment, and that this in turn affects the work pressure and the work environment and consequently (potentially) may affect sickness absence both directly and indirectly (selection). Our data allow us to investigate possible effects of children's parental background on teacher sickness by introducing indicators expressing the family background of the children in the institution.

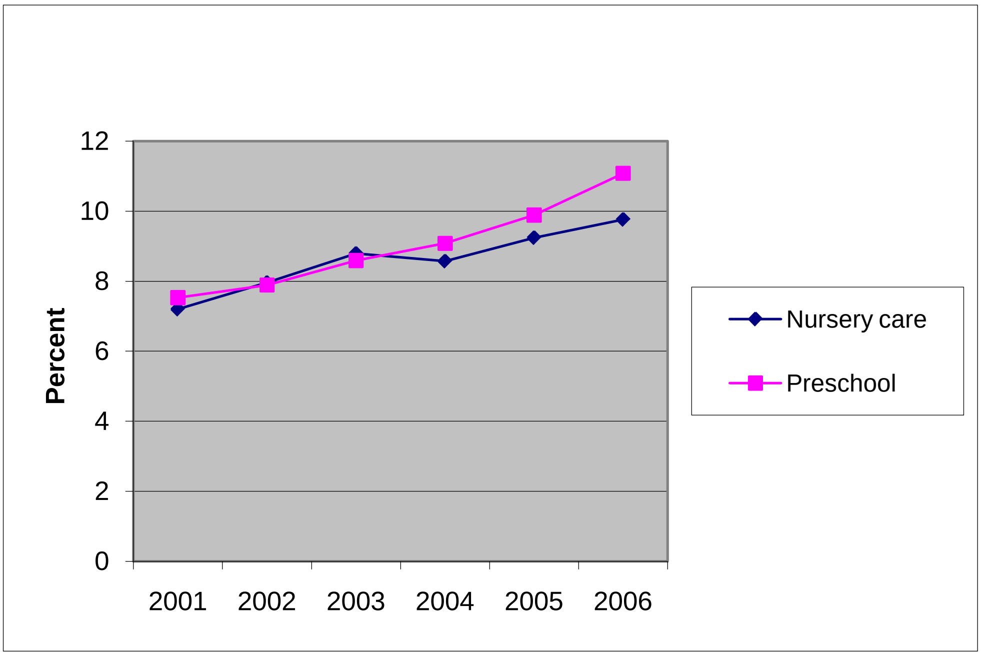
6 Sickness Absence

For the employees in the day-care sector, we distinguish between long-term sickness, which are sickness spells beyond two weeks, and total sickness, i.e. the sum of short-term and long-term sickness. We make this distinguishment, since long-term sickness absence has other characteristics than short-term sickness, see the discussion below.

6.1 Long-term sickness absence

In Denmark, the employer is liable of a refund of part of the employee's salary corresponding to social sickness benefits (around 2,000 Euro per month in 2008) for sickness absence periods longer than 2 weeks (from 2008: 3 weeks). The employer can ask for a refund already from the first day, with permission from the employee. The employer can request a medical doctor's certificate (sick note) from the employee's GP stating that the employee is indeed ill. The incidence of long-term sickness has been increasing since 2000, particularly for women. This development has been parallel to an almost constant reduction in the unemployment rate since 1993 to 4.5% of the labour force in 2006. The development in long-term sickness in the Danish day-care sector mirrors the national development. Figure 6.1 shows that the share of day-care employees who were affected by long-term sickness (for more than fifteen days) sometime during a year increased during the period 2000-2006 for nursery care and pre-school.

Figure 6.1 Share of employees affected by long-term sickness



Data source: Register data set.

6.2 Short-term sickness absence

The majority of employees (more than 9 out of 10) are not affected by long-term sickness absence at all in a given year. Since 2005, there has been a systematic account of short-term sickness absence through the municipalities' salary accounts.[3] Based on this register information, the average number of short-term sickness was around 50-55 hours for nursery care and around 46-60 hours for preschool teachers. These numbers encompass some, but not all the very long-term sickness periods.

[3] Statistics Denmark kindly made this data available for this study. To our knowledge, this analysis is the first to use individual level information on short-term sickness absence.

7 Empirical Strategy

7.1 Empirical model

According to Grossman's model of health capital, time lost due to illness reflects a depreciation of health capital (see section 3). In Grossman's theoretical formulation of the model, time lost encompasses both time lost in the labour market and time lost for non-market productive activities, i.e. home production and leisure (time at home "produces" leisure). Our empirical model of sickness absenteeism (in market work) below finds its inspiration in the formulation of the "demand" for time lost. Hence, our empirical investigation focuses on time lost in the market, while we have no account of time lost for home production and leisure, although it seems fair to assume that illness also reduces the utility of time at home.

We hypothesise that teacher absence due to sickness is a function of municipal level conditions (average child-to-teacher ratio and size of institution), workplace/institution characteristics (composition of child group) and individual characteristics (age, gender, education, family situation etc. and local unemployment rate in the municipality of residence) of the teacher. In the empirical model, the dependent variable y_{ijkt} represents teacher absence (measured by number of days or number of hours) due to sickness for teacher *i* in workplace *j* in municipality *k*, observed in period *t*:

$$y_{ijkt} = f(Z_{i,t-1}, S_{j,t-1}, \Pi_{k,t-1}, W_{it-1}, \eta_i, \phi_t, \varepsilon_{ijkt})$$

where $Z_{i,t-1}$ represents individual characteristics of teacher *i* in period *t-1*. These individual characteristics include education, age and family background of employee. $Z_{i,t-1}$ also encompasses local unemployment rate in the municipality of residence. $S_{j,t-1}$ represents firm/workplace characteristics for teacher *i* in firm *j* in period *t*. Firm characteristics may include size of firm measured by number of employees, proportion of educated pedagogical personnel in institution, and characteristics of the children (parental characteristics of the children) in the institution or municipality. $\Pi_{k,t-1}$ is average child-to-teacher ratio at the municipal level $W_{i,t-1}$ is the individual hourly wage rate.

The model reflects that we suspect that there is unobserved heterogeneity, η_i, at the individual level in for instance teacher characteristics which may affect individual sickness absence (the teacher fixed effect). η_i reflects unobserved health characteristics, social competencies, dedication, ability and the individual tendency to report absent when feeling ill. The relationship between some "objective" measure of health status and absence due to sickness is not trivial. The inclination to report absent when feeling sick varies across individuals, across firms and across sectors. Some individuals may be more inclined to report absence due to illness than others. Others may choose to go to work despite their illness, with a possible reduced productivity – this phenomenon is known as "presenteeism".

The inclination to report ill may very well be related to specific workplace conditions, individual compensation rates during illness, or depend on regional or sectoral unemployment. Social norms and attitudes towards absenteeism ("absenteeism culture") may vary across institutions, and such peer effects from colleagues may direct individuals' general perception and acceptance of absence due to sickness. In general, peer effects (from colleagues) at the workplace are known to affect individual outcomes, see e.g. Mohnen et al. (2008). Also, institutions and municipalities may have different policies towards long (or short) periods of absence which may affect the personal "choice" to report sick. Moreover, institutions and municipalities may have different sickness policies, i.e. different practices as to how soon to call in employees for a meeting with the manager after a period of long-term or repeated short-term sickness. Institutions and municipalities may also have varying policies towards firing employees who have been ill for a longer period, and this may affect the statistics on long-term illness. Thus, there may also be unobserved factors affecting individual absenteeism at the firm/institution level and at the municipal level. Unobserved heterogeneity in firm/institution characteristics (day-care institution fixed effects), which may potentially affect the outcome for teacher i, stems from peer effects from colleagues regarding attitudes towards absenteeism or heterogeneity in e.g. outdoor facilities, number of square metres and quality of management, unobserved characteristics of children's family background in the institution, proximity to transport opportunities etc. Unobserved heterogeneity in municipal characteristics (municipality fixed effects) can be due to unobserved differences in the employee policy of the municipalities, e.g. municipal policy towards sickness among employees, stated visions for child development in the day-care area etc. Whenever computationally possible, we control for the municipality fixed effect using (time-invariant) municipality dummies.

Absenteeism varies over the business cycle, as workers may be less worried about losing their job if it is relatively easy to find a job, cf. Shapiro and Stiglitz (1984). We control for aggregate time-variant shocks and time trends by using time dummies, $\phi_t.\varepsilon_{ijkt}$ is white noise.

We suspect that the unobserved individual specific effect in sickness absence is correlated with some of the explanatory variables. In particular, we hypothesise that individuals select into municipalities with favourable working conditions (including child-to-teacher ratio), and that municipalities use the child-to-teacher ratio to attract good workers (with an expected low absenteeism). Another possible type of endogeneity would occur if municipalities respond to employee sickness by reducing the child-to-teacher ratio. We examine this problem later.

7.2 Estimation techniques

In order to take unobserved heterogeneity at the individual levels into account, we take advantage of the longitudinal dimension of the data through panel data techniques. One way of taking care of this source of endogeneity is to allow for correlation between the unobserved fixed individual heterogeneity and the explanatory variables (including the child-to-teacher ratio) using panel data estimation techniques that allow for fixed effects.

Instrumental variables estimation offers an alternative way of dealing with problems of selectivity and endogeneity. As instruments for the child-to-teacher ratio we have tested two possible instruments: a dummy for whether municipalities live up to their obligation to offer day-care arrangements by the age of 12 months and the share of Conservatives/Liberals in the city council. Both instruments are lagged one year compared to the child-to-teacher ratio.

Guaranteed access to childcare (GAPS) was part of government policy from the mid-1990s, but the policy has been transmitted to municipalities in a varying pace. By 2000, around 75 per cent of the municipalities offered GAPS, and by 2006 almost 100 per cent of the municipalities report guaranteed access to childcare.[4] Our initial suspicion was that GAPS was offered more often and earlier on in municipalities with a high service level in day care and hence with a low child-to-teacher ratio. If this was the case, GAPS might suffer from the same sort of endogeneity problems as the child-to-teacher ratio. However, there have been some exogenous changes in the incentive structure set by the government in order to speed up the implementation of GAPS in the municipalities during our period of analysis. In 1999, the Danish Parliament passed a law saying that the municipalities could raise the parent user fee in day care from maximum 30 per cent of the operating expenses to 32 per cent in 2001 and 33 per cent in 2002 under the condition that municipalities offered GAPS.[5] This led around half of the municipalities to raise the parent user fee to 32 per cent in 2001, while the other half either stayed on 30 per cent or raised the fee to 31 per cent of the operating expenses. Furthermore, in 2002, 190 (out of 275) municipalities raised the parent user fee to 33 per cent of the total operating expenses in day care, cf. Bureau 2000 (2001; 2002; 2007). We expect that the decision of the municipalities to live up to GAPS affects the child-to-teacher ratio. At the same time, we find it plausible that the decision of the municipalities to implement GAPS is determined by the government incentives regarding parental fees rather than the municipality's success in attracting new day-care personnel. Thus, as we show in section 8, we have access to plausible exogenous variation in the child-to-teacher ratio. The GAPS instrument has previously been used in Datta Gupta & Simonsen (2010).

Our second instrument is the share of Conservative-Liberal politicians in the municipality city council.[6] We hypothesise (and show in section 8) that there is a positive correlation between the dummy for Conservative-Liberal majority and the child-to-teacher ratio. Moreover, we argue that this is a valid instrument as the political colour of the majority affects sickness absence and other labour-market outcomes of the day-care teachers only through the child-to-teacher ratio (or other factors which are highly correlated with the child-to-teacher ratio as for instance monetary resources to the day-care area). The share of municipalities with a Conservative-Liberal majority was 52 percent around 2000, whereas around 2 out of 3 municipalities in 2006 had a Conservative-Liberal majority. The political constitution of the municipal council may depend on the social background of the families living in

4 Information on GAPS by municipality and by year has been obtained from Niels Glavind, Bureau 2000.

5 Guaranteed access to childcare (GAPS) is fulfilled if the municipality can offer some form of childcare to families when their child reaches the age of six months. Under this guarantee, the parents cannot decide themselves whether their child should be placed in an institution rather than child minding, and the parents cannot require a specific institution in a specific (local) area.

6 The Conservative-Liberal coalition that we define here constitutes of three parties, the Conservative Party, the Liberal Party (Venstre), and Danish People's Party ("*Dansk Folkeparti*").

the municipality. We control for social background of the children in some of the regressions below.

Thus, we hypothesise (and document in section 8) that both instruments are strongly correlated with the child-to-teacher ratio and are assumed not to affect sickness absence directly, but only indirectly through the child-to-teacher ratio.

7.3 Probability of long-term absence

Our empirical strategy is motivated by the empirical distribution of sickness absence. More than 90 per cent of the workforce is not affected by long-term absence in a given year. Long-term absence (more than two weeks) is to a certain extent a 1/0 decision. Therefore, we initially analyse the probability of becoming long-term ill.

The model for the probability of long-term sickness absence follows the standard specification of binary response models with longitudinal data. We model an underlying latent variable y_{ijkt}^{*}:

$$y_{ijkt}^{*} = \beta_0 + \beta_1 Z_{i,t-1} + \beta_2 S_{j,t-1} + \beta_3 \Pi_{k,t-1} + \beta_4 W_{i,t-1} + \eta_i + \phi_t + \varepsilon_{ijkt} \quad (1)$$

$$where\ y_{ijkt} = 1[y_{ijkt}^{*} > 0]$$

Thus, $y=1$ if an employee has been long-term ill in year t and $y=0$ if he/she has not been long-term ill. We model the conditional probability of long-term illness ($y=1$) using the logit function which, in a panel context, allows for correlation between unobserved heterogeneity and the explanatory variables in a fixed effects framework.

As an alternative, we also experiment with the linear probability model, where we instrument the child-to-teacher ratio using the two instruments, GAPS and the dummy for whether the municipality has a majority of Conservatives and Liberals, controlling for municipality fixed effects by using municipality dummies.

7.4 Number of days in short- and long-term absence

The duration of long-term sickness spells varies from 0-360 days a year. We analyse the number of days being absent due to illness using the fixed effects Poisson estimator which was originally developed by Hausman, Hall & Griliches (1984). The Poisson estimator is especially well suited for analyses of count variables as the number of days absent from work, cf. Wooldridge (2003). In this empirical specification of the model, the length of sickness absence (measured in days), y, is modelled as $y*$ in the model above.

$$y_{ijkt} = \beta_0 + \beta_1 Z_{it-1} + \beta_2 S_{jt-1} + \beta_3 \Pi_{kt-1} + \beta_4 W_{it-1} + \eta_i + \phi_t + \varepsilon_{ijkt} \quad (2)$$

The density of y given observables under the Poisson assumption is completely determined by the conditional mean. Thus, the Poisson distributional assumption imposes a number of restrictions on the conditional moments of y. The most important restriction, which is often violated, is that there is equality of the conditional variance and the conditional mean: $var(y|X)=E(y|X)$, where X encompasses all explanatory and control variables – the so-called Poisson variance assumption. A weaker assumption allows the variance-mean ratio to be any positive constant. When this ratio is larger than 1, which is the case with our data, the situation is called overdispersion. In cases of overdispersion, the literature usually suggests using a particular form of the Poisson model, i.e. the NegBin (II) model, cf. Cameron & Trivedi (2009). Count data models based on longitudinal data usually assume that the unobserved effect has a multiplicative form. This form allows for an arbitrary dependence between the unobserved effect and the explanatory variables, cf. Wooldridge (2003).

As an alternative, we also analyse number of days in long-term absence in a linear framework (linear probability model, LPM) where we can instrument the child-to-teacher ratio and control for municipality level fixed effects using municipality dummies.

Periods of short-term absence are much more prevalent, and most employees have a positive number of sickness hours in a year. The number of hours of short-term sickness absence is analysed by both the fixed-effects Negbin model and in a linear framework where the child-to-teacher ratio is instrumented using the instruments mentioned in section 7.2.

8 Empirical Results

The empirical analysis estimates models (1) and (2). We analyze the determinants of sickness absence and possible states after a long-term sickness absence period using the following dependent variables:

1 *Probability of long-term absence* is analysed by using the fixed-effects logit estimator (model (1)) and the linear probability model with instruments for the child-to-teacher ratio.

2 *Number of days of long-term absence* is analysed by using the fixed-effects negative binominal estimator and linear instrumental variables (2SLS) regression (model (2)).

3 *Number of days of short-term absence* is analysed by using the fixed-effects negative binomial estimator and (2SLS) instrumental variables estimation (model (2)).

Regressions are performed separately for nursery care and preschool. Our focus is on the child-to-teacher ratio, and we control for individual characteristics, the local unemployment rate. We include year dummies in all models and municipality dummies in the linear versions of the model, i.e. the LPM model and the 2SLS regression.

Our primary parameter of interest is the child-to-teacher ratio. We analyse the relationship between long-term sickness absence and one-year-lagged child-to-teacher ratio in order to avoid that a possible correlation might in fact be due to action taken by the municipalities to accommodate a high level of sickness absence by adjusting child-to-teacher ratios (reversed causality). Thus, long-term sickness absence is analysed separately for 2002-2004 (using child-to-teacher ratios from 2001-2003) and 2005-2006 (using child-to-teacher ratios from 2004-2005). The two periods are analysed separately due to the change in the definition of child-to-teacher ratios from 2003-2004, see section 5 for a discussion. We analyse total sickness absence for the later period (2005-2006, where information on short-term sickness absence is available). For the estimations on short-term absence, we use the child-to-teacher ratio in the same year, as we expect that especially short-term sickness periods are mainly affected by the child-to-teacher ratio in the same year.

Tables with regression results are shown at the end of the paper. Table 1 shows summary statistics of the sample used in the estimations of long-term sickness. The child-to-teacher ratio is measured by the average number of children per full-time pedagogical employee at the municipal level. Alternatively, one might consider using the number of children per trained teacher (*pædagog*) as a measure of the child-to-teacher ratio. The correlation between this alternative measure and the simple measure used here is more than 0.9, and a sensitivity analysis using this alternative measure of the child-to-teacher ratio did not change the results significantly.

8.1 Probability of long-term sickness absence

The results of the *first* step of the empirical analyses, i.e. the fixed-effects logit estimations and linear IV estimations of the probability of long-term sickness absence (more than two weeks a year), are shown in tables 2-3. The Hausman test rejects that fixed-effects and random-effects results are not significantly different. Hence, fixed-effects estimation is preferred. The FE model is potentially unstable due to few time periods. Moreover, the number of observations is reduced substantially since FE logit drops observations with all positive or all negative outcomes and individuals that are only observed in one period. Hence identification is obtained from individuals that change status within the period. For nursery care (table 2), there is no significant relationship between the child-to-teacher ratio and the probability of experiencing a long-term absence period in a year for neither the FE logit nor the linear probability IV models. For preschool teachers (table 3), we find a positive and significant relationship between the child-to-teacher ratio and the probability of long-term sickness absence when focusing on 2005-06 (model 5). This implies that an increase in the child-to-teacher ratio by one child increases the average probability of becoming long-term absent with 4 percentage points (compared to an average probability of long-term sickness absence of 10% for this group). This effect becomes statistically insignificant when controlling for institution characteristics. Instead, the share of immigrant parents among the children seems to be positively related to the probability of long-term sickness absence.

The net hourly wage has a negative and significant effect on the probability of long-term sickness absence in nursery, as predicted by Grossman's health model. Age and age squared are significant, and experience, which is highly correlated with age, has a negative and significant effect on the risk of becoming long-term ill for preschool teachers. The number of children below 18 has a negative effect on the incidence of long-term sickness periods. Since we also explicitly control for age and tenure, this is not an age effect.[7]

8.2 Length of long-term absence

In our *second empirical step*, we analyse the determinants of the length of long-term sickness periods using two different estimation techniques; the fixed-effects Negbin estimator and instrumental variables estimation. The FE Negbin results are shown in table 4 (nursery care) and table 5 (preschool). FE Negbin identifies the effects based on individuals who change status and who are observed for at least two periods. Due to the majority of the employees having no long-term periods of sickness absence, identification is based on a reduced sample

[7] Paringer (1983) found that the presence of family responsibilities appears to reduce the amount of time missed from work, particularly among women, and she suggests that an explanation of this might be that the dual responsibilities of women with a family might induce them to invest more in their health thereby lowering their illness rates. Paringer also suggests that because of women's dual responsibilities, the full impact of ill health may carry with it greater costs than the lost earnings associated with missing work. Thus, women may have a lower work loss threshold to a given illness than men because there is a greater payoff to the family if the woman responds to the illness earlier on. Another explanation might be common causality: that factors that determine the likelihood of marriage and fertility also affect later health outcomes and the propensity to report absent from work over a given health condition.

of observations with these characteristics. This reduces significance. Hence, the number of observations using fixed-effects Negbin is reduced substantially compared to the original data, since this estimation method drops observations with only one observation per group and furthermore drops groups of observations (individuals) with all zero outcomes.

FE Negbin identifies the effects based on individuals who change status. Due to the majority of employees with no long-term sickness absence periods, identification is based on individuals that change status, and this reduces significance of the estimated parameters. For both time periods, we find no significant relationship between the number of days in long-term sickness and (lagged) child-to-teacher ratio when using the FE Negbin estimator. We now turn to the instrumental variables (IV) results. The child-to-teacher ratio is instrumented by the two instruments mentioned above, the guaranteed access to childcare (GAPS) and a dummy for whether the political majority in the municipality is Conservative-Liberal. First-stage results (see Appendix 1) reveal that both instruments are strongly correlated with the child-to-teacher ratio. Moreover, the overidentifying restrictions (Sargan) test does not reject that our two instruments are exogenous given that at least one instrument is exogenous. Under this assumption, the two instruments are valid instruments.

Turning now to the second-stage results of the IV-estimation in table 4-5, we find that the relationship between the child-to-teacher ratio and the length of sickness absence is insignificant. Thus, although we find a positive and significant relationship between the child-to-teacher ratio and the *probability* of becoming long-term ill, the *length* of absence periods are not significantly related to the child-to-teacher ratio once we take care of selectivity. This is perhaps not surprising: While the child-to-teacher ratio may affect the probability of exit to long-term sickness absence, once a person is long-term ill, many other factors will be important for the return to work, including the sickness policy of the employer, whether employees on long-term sickness absence tend to be laid off after some time, options for flexible return to work etc.

For the remaining explanatory variables, we get similar findings as for the model for the probability of long term sickness absence discussed above. Experience is negatively related to length of long-term sickness absence, and the local unemployment rate has a negative effect, as expected. Moreover, workplace characteristics play a significant role for the length of preschool sickness absence. The share of trained teachers reduces length of long-term absence, while the share of immigrant children seems to be associated with a higher extent of sickness absence among the personnel.

Another endogeneity problem stems from the suspicion that municipalities or day-care institutions choose the level of child-to-teacher ratio based on the experienced level of sickness absence in the previous year. If sickness absence in period t is correlated with lagged sickness absence in nursery care and preschool, there is a risk that child-to-teacher ratio in period t is decided based on the average municipality absence due to sickness in t-1. We investigated this by regressing municipality changes in the child-to-teacher ratio on lagged changes in average long-term absenteeism at the municipal level. However, we found no significant correlation here, so there is no statistically significant indication that municipalities try to mitigate a high level of absence by devoting more personnel resources to day care.

8.3 Length of short-term sickness absence

In the *third step*, we analyse the relationship between the child-to-teacher ratio and the length of short-term sickness absence. While only around 15 per cent of the sample experience a long-term absence period during a year, the incidence of short-term periods of absence is much more widespread.

Again, we employ FE Negbin as well as IV estimation using the same instruments as above. Estimation results show that there is a positive relationship between the child-to-teacher ratio and short-term absence due to sickness for nursery-care teachers, cf. table 6, columns 1 and 3. However, the relationship between the child-to-teacher ratio and short-term sickness is insignificant for preschool, cf. table 7. For nursery care, the IV results in model 3 indicate that an increase in the child-to-teacher ratio by 1 child per teacher (which is a relatively large change in the child-to-teacher ratio for nursery care) is associated with 16 hours more sickness absence per year. Compared to an average sickness absence of 57 hours per year for nursery-care teachers, this amounts to an increase in short-term absence of almost 30%. We find no significant effects of institution size or other institution characteristics, and the results of the regressions controlling for institution characteristics are not shown. Individual characteristics have a strong and significant effect on short-term absence. Age, experience and being a trained teacher are positively related to sickness absence, while women teachers have a lower level of absence. Number of children and hourly wage are negatively correlated with sickness absence. Somewhat counterintuitive, the unemployment rate is positively related to sickness absence.

8.4 Discussion

The child-to-teacher ratio has a positive and significant impact on sickness absence for 2005-2006, while we find no significant effects before 2004. Hence, although there is a somewhat clear statistical relationship for some periods, the evidence is somewhat mixed. Unobserved factors at the municipality level (municipality dummies) usually strong and significant impact implying that other factors at the municipality level are also potentially important. These may include the employer's sickness absence policy, options for flexible return to work, firing practices after a sickness absence period etc. These practices have been developed considerably over the last 1-2 decades. Since we find substantial variation in levels of sickness across institutions in a municipality, we need more details. Other factors that are potentially important are the options for professional development, the quality of management etc. These observations are consistent with the results in Lindeboom and Kerkhofts (2000).

8.5 Public finance simulation

The estimation results indicate that for some specifications of the model, both short- and long-term sickness absence is rising with the child-to-teacher ratio. For example, we find in section 8.2 that an increase in the child-to-teacher ratio of 1 child per employee is associated with an increase in the risk of becoming long-term sick (i.e. absent for more than 2 weeks) of

4 percentage points or 40% compared to the average risk of long-term absence of 10% in preschool.

For an average kindergarten with around 60 children and a child-to-teacher ratio of 5.75 children per teacher, an increase by 1 child per teacher leads to a reduction in the staff by 1 teacher or a reduction in the the labour costs of around 17%. The average yearly salary for teachers was 360,000 DKR in the beginning of 2010.

However, an increase in the child-to-teacher ratio of 1 leads to an increase in sickness absence of 0.04 or 40% of the average frequency of long-term absence of 10%. With a staff of 10 employees, there was previously one employee long-term ill per year, so the increase in employees on long-term sickness leave is 0.4. The average duration of a period of long-term absence is 120 days or 1/3 of a year. Thus, if sickness absence periods are replaced by substitute teachers, yearly labour costs for substitute teachers increase by 44,000 DKK.

Moreover, the risk of exiting the labour market and going on disability pension from a long-term sickness period is 4.9 percent in our sample. Thus, saving one teacher leads to an increase in disability pension risk of some 2% (4.9%*0.4). The average age of a person going on disability pension is 51, which is more than 10 years before average retirement age, and the average yearly rate for disability pension was around 180,000 DKK in 2010. Thus, saving one teacher's salary one year leads to extra disability costs of 35,000 DKK (0.02*10*180,000 DKK).

All in all, saving one teacher salary leads to extra expected expenses to sickness absence replacements and disability pensions (of around 44,000+35,000=79,000 DKK). The extra costs amount to around 22% of the saved costs.

The calculation above only includes increases in long-term sickness absence. Moreover, fewer teacher resources may impair the quality of day-care offerings, lead to higher sickness among the children, and hence ultimately costs to the parents in terms of foregone earnings and reduced productivity. This is not included in the calculation above.

9 Conclusion

This paper analyses various aspects of sickness absence among Danish day-care teachers. We are particularly interested in studying the role of work pressure as measured by the child-to-teacher ratio. Teachers may select into institutions/municipalities with favourable work conditions, including generous child-to-teacher ratio. Thus, a possible correlation between child-to-teacher ratio and outcome variables is not necessarily causal. We consider various modes of dealing with this endogeneity, i.e. exploiting the panel dimension of the data and instrumental variable methods. We employ two instruments: a dummy for whether municipalities provide guaranteed access to childcare (GAPS) and the share of Conservative-Liberal local politicians in the municipal city council. These instruments are shown to be relevant (correlated with our endogenous explanatory variable) and pass the overidentification test whereby we can establish that the instruments are exogenous if at least one instrument is exogenous.

First, we look at the probability of becoming long-term ill, estimated by fixed-effects logit and 2SLS. Our estimation results indicate that the probability of becoming long-term ill is positively related to the child-to-teacher ratio for preschool in 2005-2006. So work pressure enhances the risk of becoming long-term ill for preschool teachers in the latter period, but not enough for 2002-2004

Secondly, we analyse the length of absence periods, but we find no significant relationship to the child-to-teacher ratio. This may seem difficult to reconcile in our model. However, this result does indicate that once one has passed this private threshold for reporting long-term sick, other factors seem more important.

Thirdly, we look at how the extent of short-term absence is related to the child-to-teacher ratio. Our estimation results indicate that nursery-care teachers' sickness absence is positively related to their work pressure, as suggested by our theoretical model. However, we find no significant relationship for preschool teachers.

Hence, all in all, our estimation results are mixed, although we do get an indication of a positive and significant relationship. Factors like management, organisation, sickness policy, in generation seem to be important in assessing the background of high child-to-teacher ratio.

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Appendix 1: Validity of Instruments

We take a look at the first-stage regressions in the IV estimation. Our instruments should fulfil two criteria: Relevance and exogeneity.

First, instruments should be relevant in explaining the variation in the child-to-teacher ratio. We find that instruments, i.e. the municipality guaranteed child-care provision (lagged) and the political colour of the city council, are strongly correlated with the child-to-teacher ratio. A rule of thumb is that the F-statistics for the test of the joint significance of the instruments in the first-stage regression should be above 10. This criterion is fulfilled for nursery care before 2004, cf. table A1.1 for first-stage regression results for length of long-term absence (tables 6-7). Guaranteed access to child care is generally negatively related to the child-to-teacher ratio. Thus, municipal politicians' guarantee of child-care access was presumably followed up by more resources. The period 2005-2006 for preschool is an exception to this observation. Moreover, the presence of a majority of Conservative-Liberals in the city council seems to be associated with a higher child-to-teacher ratio (and fewer resources to child care).

Table A1.1: First-stage regressions for long-term absence

	Nursery care			
	2002-2004		2005-2006	
	b	t	b	t
GAPS	-0.270	5.02	-0.074	32.52
Share Conservative-Liberals	0.002	0.18	0.078	8.56
F-value*	0.556		3.23	

*) F-value for test of joint significance of the two instruments.

Note: The first-stage regression corresponds to IV regressions in table 6, column 1 and 3. First-stage regression also includes exogenous explanatory variables included in the main model.

Table A1.2: First-stage regressions for long-term absence

	Preschool			
	2002-2004		2005-2006	
	B	t	b	t
GAPS	-0.119	13.49	0.229	49.06
Share Conservative-Liberals	0.027	1.48	-0.146	8.43
F-value*	2.20		15.83	

*) F-value for test of joint significance of the two instruments.

Note: The first-stage regression corresponds to IV regressions in table 7, column 1 and 3. First-stage regression also includes exogenous explanatory variables included in the main model.

Secondly, the instruments should be exogenous to the unobserved variation in our sickness absence measure. This is tested by the Sargan test, which is a test of the overidentifying restrictions. The hypothesis being tested with the Sargan test is that the instrumental variables are uncorrelated to some set of residuals given that at least one instrument is exogenous. In that case, they are acceptable, healthy instruments. We find that the null hypothesis is accepted statistically in all the IV estimations in section 8.2 and 8.3, so the instruments pass the overidentifying restrictions test and hence are valid by this criterion.

Tables

Table 1: Summary statistics

	Nursery care				Preschool			
	Mean	Std.	Min	Max	Mean	Std.	Min	Max
Long-term sick days	11.009	45.259	0.000	360.000	10.747	43.899	0.000	360.000
Short-term sickness hours	55.729	46.658	0.000	814.000	47.992	41.303	0.000	1359.425
Child-to-teacher ratio	2.818	0.260	1.440	4.830	5.748	0.821	2.140	10.390
Age	38.911	12.098	15.000	65.000	40.894	11.426	14.000	65.000
Number of children	0.683	0.942	0.000	6.000	0.823	1.020	0.000	7.000
Experience	13.617	10.229	0.000	43.000	14.701	9.628	0.000	43.000
Woman	0.940	0.238	0.000	1.000	0.889	0.314	0.000	1.000
Trained teacher	0.384	0.486	0.000	1.000	0.439	0.496	0.000	1.000
Single	0.289	0.453	0.000	1.000	0.214	0.410	0.000	1.000
ln net hourly wage	4.098	0.332	-0.777	5.285	4.124	0.342	-0.806	5.293
Unemp rate	5.704	1.397	2.200	12.700	5.794	1.629	2.200	15.900
Size of institution	13.260	1.730	2.000	19.900	8.537	1.662	2.500	20.700
Share of trained teacher	0.527	0.143	0.000	1.000	0.581	0.162	0.000	1.000
Share of male teachers	0.098	0.084	0.000	1.000	0.123	0.093	0.000	1.000
Share of immigrant parents	0.130	0.129	0.000	0.944	0.095	0.114	0.000	0.950
Share of parents with short education	0.194	0.049	0.071	0.428	0.203	0.052	0.011	0.456
N	43418				130891			

*) Short-term sickness absence is only observed in 2005-2006.

Table 2: Probability of long-term* sickness absence for nursery-care teachers

	2002-2004			2005-2006		
	FE Logit	LPM, IV	LPM, IV	FE Logit	LPM, IV	LPM, IV
	b/t	b/t	b/t	b/t	b/t	b/t
Child-to-teacher ratio	-0.140	-0.639	-0.745	0.266	-0.225	-0.545
	(0.50)	(1.28)	(1.52)	(0.53)	(1.77)	(1.27)
Personal characteristics						
Woman		0.030***	0.029***		0.027***	0.028***
		(5.31)	(5.73)		(4.39)	(4.21)
Teacher	-0.001	-0.004	-0.001	0.021	-0.006	-0.005
	(0.01)	(0.78)	(0.16)	(0.07)	(1.15)	(1.04)
Ln age	-228.602*	0.733***	0.752***	-320.213	0.165	0.132
	(2.04)	(3.42)	(3.37)	(1.56)	(0.60)	(0.45)
Ln age sq	44.178*	-0.090**	-0.092**	65.101	-0.008	-0.004
	(2.09)	(3.03)	(3.01)	(1.69)	(0.22)	(0.09)
Dummy single	-0.216	-0.005	-0.005	0.662	-0.005	-0.005
	(0.79)	(0.96)	(1.03)	(1.45)	(1.00)	(0.95)
# children	0.139	-0.011**	-0.011**	-0.859	-0.000	0.001
	(0.45)	(3.02)	(2.82)	(1.53)	(0.02)	(0.11)
Experience	0.448	-0.000	0.000	0.925	0.005	0.005
	(0.51)	(0.09)	(0.00)	(0.56)	(0.96)	(0.94)
laglnwagenet	-0.600	0.011	0.011	-1.354*	0.010	0.010
	(1.94)	(1.63)	(1.62)	(2.42)	(1.33)	(1.36)
Unemp rate	-0.001	0.004	0.005	-0.054	-0.004	-0.004
	(0.01)	(0.79)	(1.10)	(0.39)	(1.72)	(1.56)
Institution characteristics						
Size of inst			-0.077			-0.088
			(1.26)			(1.17)
Sh educ emp			-0.070			-0.015
			(1.45)			(1.12)
Sh male emp			0.021			0.021
			(1.13)			(0.86)
Sh immigr			0.037			-0.013
			(1.92)			(0.67)
Sh no educ			0.347			0.110
			(0.96)			(0.45)
Constant		0.183	1.518		0.243	2.416
		(0.13)	(0.77)		(0.33)	(0.91)
Municipality dummies	No	Yes	Yes	No	Yes	Yes
Time dummies	Yes	Yes	Yes	Yes	Yes	Yes
Log likelihood	-1192			-480		
N	3388	23835	23807	1418	15119	15022

*) Long-term: More than 15 days in a year.

t-values in parentheses. Standard errors are municipality clustered. *) p=0.05, **) p=0.01, ***) p=0.001

Table 3: Probability of long-term* sickness absence for preschool teachers

	2002-2004			2005-2006		
	FE Logit	LPM, IV	LPM, IV	FE Logit	LPM, IV	LPM, IV
	b/t	b/t	b/t	b/t	b/t	b/t
Child-to-teacher ratio	-0.029	0.029	0.041	0.142	0.040*	0.042
	(0.61)	(0.58)	(0.49)	(1.57)	(1.99)	(1.58)
Personal characteristics						
Woman		0.012***	0.012***		0.022***	0.022***
		(3.66)	(3.71)		(3.95)	(4.07)
Teacher	0.097	-0.002	-0.002	0.521**	-0.010**	-0.010**
	(0.73)	(0.95)	(0.90)	(3.25)	(3.13)	(3.03)
Ln age	-154.366*	0.772***	0.787***	-386.142**	1.142***	1.140***
	(2.17)	(6.38)	(6.48)	(2.96)	(8.22)	(8.05)
Ln age sq	28.731*	-0.093***	-0.095***	73.891**	-0.142***	-0.141***
	(2.16)	(5.66)	(5.75)	(3.05)	(7.54)	(7.39)
Dummy single	-0.063	-0.004	-0.004	0.048	0.008	0.009
	(0.38)	(1.19)	(1.21)	(0.19)	(1.77)	(1.82)
# children	0.135	-0.005	-0.005	0.193	-0.009*	-0.010**
	(0.83)	(1.67)	(1.70)	(0.76)	(2.51)	(2.70)
Experience	1.494*	-0.010**	-0.011**	1.045	-0.010*	-0.010*
	(2.48)	(2.86)	(3.05)	(0.99)	(2.39)	(2.30)
laglnwagenet	-0.178	-0.006	-0.005	-0.450	-0.005	-0.004
	(1.09)	(1.43)	(1.31)	(1.49)	(0.90)	(0.72)
Unemp rate	-0.081	-0.003	-0.003	-0.170*	-0.004*	-0.004*
	(1.51)	(1.50)	(1.50)	(1.99)	(2.21)	(2.00)
Institution characteristics						
Size of inst			0.019			0.009
			(0.49)			(0.91)
Sh educ emp			-0.008			-0.004
			(1.01)			(0.39)
Sh male emp			0.002			0.007
			(0.12)			(0.41)
Sh immigr			0.011			0.034*
			(1.56)			(2.28)
Sh no educ			0.014			-0.091
			(0.24)			(1.55)
Constant		-1.574***	-1.831*		-2.315***	-2.389***
		(4.90)	(2.39)		(8.80)	(7.35)
Municipality dummies	No	Yes	Yes	No	Yes	Yes
Time dummies	Yes	Yes	Yes	Yes	Yes	Yes
Log likelihood	-3534			-1608		
N	10051	71686	71109	4768	46244	45774

*) Long-term: More than 15 days in a year.

t-values in parentheses. Standard errors are municipality clustered. *) p=0.05, **) p=0.01, ***) p=0.001

Table 4: Number of long-term sickness days, nursery care

	2002-2004				2005-2006			
	FE Negbin	FE Negbin	IV	IV	FE Negbin	FE Negbin	IV	IV
	b/t	b/t	b/t	b/t	b/t	b/t	b/t	b/t
Child-to-teacher ratio	0.042	0.010	-12.751	-45.637	-0.260	-0.283	-9.563	-37.325
	(0.34)	(0.07)	(0.24)	(1.14)	(1.45)	(1.22)	(0.87)	(1.11)
Personal characteristics								
Woman	0.403	0.405	4.226***	4.383***	0.804*	0.757*	3.561***	3.704***
	(1.51)	(1.50)	(5.27)	(5.85)	(2.16)	(2.03)	(5.26)	(5.24)
Teacher	-0.119	-0.115	0.544	0.738	-0.168	-0.189	0.220	0.255
	(1.66)	(1.58)	(0.93)	(1.26)	(1.75)	(1.95)	(0.26)	(0.30)
Ln age	10.665*	10.761*	53.407*	53.522*	18.251**	18.118**	17.006	15.024
	(2.33)	(2.34)	(1.99)	(1.97)	(2.71)	(2.69)	(0.44)	(0.39)
Ln age sq	-1.362*	-1.371*	-6043.000	-6084.000	-2.360**	-2.339*	-0.433	-0.165
	(2.19)	(2.19)	(1.59)	(1.60)	(2.60)	(2.57)	(0.08)	(0.03)
Dummy single	0.020	0.020	-0.154	-0.305	-0.053	-0.047	-0.577	-0.519
	(0.23)	(0.23)	(0.26)	(0.47)	(0.41)	(0.36)	(0.57)	(0.53)
# children	-0.075	-0.074	-0.690	-0.669	-0.303*	-0.320*	0.467	0.524
	(0.86)	(0.83)	(1.42)	(1.33)	(2.34)	(2.47)	(0.49)	(0.56)
Experience	-0.305***	-0.308***	0.444	0.533	-0.222	-0.219	0.259	0.268
	(3.45)	(3.46)	(0.57)	(0.73)	(1.83)	(1.80)	(0.35)	(0.36)
laglnwagenet	-0.125	-0.126	1.355*	1.366*	-0.309	-0.298	1.498	1.474
	(1.01)	(1.01)	(2.24)	(2.16)	(1.73)	(1.66)	(1.93)	(1.91)
Unemp rate	0.043	0.048	0.546	0.680	-0.016	-0.026	-0.046	-0.066
	(1.56)	(1.64)	(1.61)	(1.88)	(0.46)	(0.73)	(0.13)	(0.18)

	2002-2004				2005-2006			
	FE Negbin	FE Negbin	IV	IV	FE Negbin	FE Negbin	IV	IV
	b/t	b/t	b/t	b/t	b/t	b/t	b/t	b/t
Institution characteristics								
Size of inst		-0.016		-4.931		-0.017		-7.984
		(0.64)		(1.08)		(0.45)		(1.32)
Sh educ emp		-0.138		-5.637		0.074		-0.381
		(0.73)		(1.47)		(0.32)		(0.25)
Sh male emp		0.235		3.997		-0.103		2.933
		(0.59)		(1.79)		(0.24)		(0.92)
Sh immigr		0.295		3.277		0.045		2.410
		(1.50)		(1.35)		(0.18)		(1.45)
Sh no educ		-0.427		37.051		0.992		2.569
		(0.59)		(1.35)		(1.08)		(0.16)
Constant	-22.833**	-22.668**	-84.699	66.441	-35.297**	-34.934**	-26.819	164.499
	(2.73)	(2.69)	(0.53)	(0.36)	(2.87)	(2.84)	(0.38)	(0.83)
Municipality dummies	No	No	Yes	Yes	No	No	Yes	Yes
Time dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Log likelihood	-7259	-7246			-3527	-3504		
N	5926	5915	23835	23807	2972	2946	15119	15022

*) Long-term: More than 15 days in a year.

t-values in parentheses.

Standard errors are municipality clustered.

*) p=0.05, **) p=0.01, ***) p=0.001

Table 5: Number of long-term* sickness days, preschool

	2002-2004				2005-2006			
	FE Negbin	FE Negbin	IV	IV	FE Negbin	FE Negbin	IV	IV
	b/t	b/t	b/t	b/t	b/t	b/t	b/t	b/t
Child-to-teacher ratio	-0.005	-0.006	4.778	8.606	-0.043	-0.043	1.574	0.428
	(0.22)	(0.24)	(0.70)	(0.71)	(1.28)	(1.13)	(0.47)	(0.10)
Personal characteristics								
Woman	-0.294***	-0.313***	1.380**	1.504***	0.114	0.082	2.807***	2.867***
	(3.39)	(3.51)	(3.19)	(3.32)	(0.93)	(0.66)	(3.58)	(3.72)
Teacher	-0.210***	-0.208***	0.001	0.004	-0.258***	-0.246***	-0.717	-0.798
	(5.21)	(5.05)	(0.00)	(0.01)	(5.03)	(4.71)	(1.46)	(1.69)
Ln age	18.516***	18.732***	92.064***	91.093***	13.956***	13.936***	123.389***	123.795***
	(6.23)	(6.26)	(6.40)	(6.31)	(3.56)	(3.52)	(8.56)	(8.48)
Ln age sq	-2.396***	-2.424***	-11.009***	-10.851***	-1.723**	-1.717**	-15.273***	-15.309***
	(5.97)	(6.00)	(5.65)	(5.54)	(3.26)	(3.22)	(7.80)	(7.70)
Dummy single	-0.042	-0.041	-0.731	-0.774*	0.042	0.038	0.805	0.874
	(0.71)	(0.71)	(1.90)	(2.05)	(0.56)	(0.50)	(1.32)	(1.45)
# children	-0.129**	-0.134**	-1.018*	-1.012*	-0.109	-0.105	-0.627	-0.689
	(2.63)	(2.72)	(2.32)	(2.28)	(1.62)	(1.55)	(1.33)	(1.44)
Experience	-0.383***	-0.400***	-1.568***	-1.610***	-0.405***	-0.418***	-0.805	-0.818
	(7.84)	(8.07)	(3.89)	(3.99)	(6.40)	(6.46)	(1.63)	(1.66)
laglnwagenet	-0.214***	-0.199***	-0.657	-0.623	-0.540***	-0.522***	-0.327	-0.209
	(3.66)	(3.37)	(1.24)	(1.15)	(5.72)	(5.44)	(0.56)	(0.34)
Unemp rate	0.014	0.018	-0.376	-0.341	0.030	0.031	-0.508*	-0.503*
	(1.04)	(1.29)	(1.27)	(1.17)	(1.85)	(1.85)	(2.47)	(2.42)
Institution characteristics								
Size of inst		0.006		4.253		-0.000		-0.442
		(0.49)		(0.74)		(0.02)		(0.28)

	2002-2004				2005-2006			
	FE Negbin	FE Negbin	IV	IV	FE Negbin	FE Negbin	IV	IV
	b/t	b/t	b/t	b/t	b/t	b/t	b/t	b/t
Sh educ emp		-0.003		-0.832		-0.278*		0.202
		(0.03)		(0.77)		(2.43)		(0.13)
Sh male emp		-0.204		1.876		-0.024		1.324
		(1.07)		(0.92)		(0.11)		(0.60)
Sh immigr		-0.279		4.209**		0.033		4.288*
		(1.88)		(3.09)		(0.21)		(2.15)
Sh no educ		-0.225		8.264		-0.386		-0.973
		(0.59)		(0.89)		(0.87)		(0.14)
Constant	-36.076***	-36.453***	-194.965***	-251.582*	-26.911***	-26.703***	-238.261***	-231.705***
	(6.64)	(6.66)	(4.55)	(2.24)	(3.74)	(3.68)	(8.14)	(5.49)
Mun. Dum	No	No	Yes	Yes	No	No	Yes	Yes
Time dum.	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Log likelihood	-2.11e+04	-2.08e+04			-1.15e+04	-1.13e+04		
N	17733	17558	71686	71109	9584	9420	46244	45774

*) Long-term: More than 15 days in a year.

t-values in parentheses.

Standard errors are municipality clustered.

*) p=0.05, **) p=0.01, ***) p=0.001

Table 6 Short-term sickness absence, nursery care, 2005-06

	(1) Negbin b/t	(2) FE Negbin b/t	(3) 2SLS b/t
Child-to-teacher ratio	0.203***	0.037	16.425***
	(5.37)	(0.77)	(8.72)
Personal characteristics			
Dummy woman	-0.069	-0.005	-2.851
	(1.22)	(0.04)	(0.96)
Dummy teacher	0.015	0.095*	0.367
	(0.53)	(2.08)	(0.22)
ln age	5.054***	8.669***	275.171***
	(5.82)	(3.54)	(7.62)
ln age squared	-0.754***	-1.203***	-40.922***
	(6.36)	(3.60)	(8.42)
Dummy single	0.096***	0.034	5.693***
	(5.57)	(0.64)	(5.38)
# children	-0.054**	0.017	-3.050*
	(2.63)	(0.31)	(2.38)
Experience	0.177***	0.160*	8.947***
	(9.20)	(2.42)	(5.65)
ln hourly wage net	-0.032	-0.351***	-0.802
	(0.71)	(4.03)	(0.46)
Unemployment rate	0.009	-0.007	0.407
	(0.95)	(0.45)	(0.86)
Constant	-5.045***	-13.467**	-456.282***
	(3.45)	(3.06)	(7.05)
ln alpha	-0.551***		
	(32.43)		
log likelihood	-4.50e+04	-1.28e+04	
N	9085	5946	9085

t-values in parentheses. Standard errors are municipality clustered.

*) p=0.05, **) p=0.01, ***) p=0.001

Table 7 Short-term sickness absence, preschool, 2005-06

	(1) Negbin b/t	(2) FE Negbin b/t	(3) 2SLS b/t
Child-to-teacher ratio	0.015	0.003	5.673
	(0.93)	(0.22)	(0.30)
Personal characteristics			
Dummy woman	-0.099***	-0.024	-4.674***
	(4.57)	(0.44)	(4.19)
Dummy teacher	0.083***	0.079***	3.750***
	(6.74)	(3.44)	(6.42)
ln age	6.273***	6.009***	304.292***
	(9.15)	(3.54)	(9.42)
ln age squared	-0.900***	-0.834***	-43.410***
	(9.55)	(3.64)	(9.80)
Dummy single	0.117***	0.041	6.061***
	(8.03)	(1.18)	(7.87)
# children	-0.052***	-0.001	-2.422***
	(4.25)	(0.05)	(4.14)
Experience	0.113***	0.183***	4.655***
	(7.12)	(4.13)	(6.73)
ln hourly wage net	-0.042	-0.273***	-1.493
	(1.68)	(5.46)	(1.45)
Unemployment rate	0.016*	0.012	0.773*
	(2.20)	(1.54)	(2.18)
Constant	-6.965***	-9.139**	-501.983***
	(5.60)	(2.96)	(4.40)
ln alpha	-0.532***		
	(55.20)		
log likelihood	-1.34e+05	-3.85e+04	
N	27822	18390	27822

t-values in parentheses. Standard errors are municipality clustered.

*) p=0.05, **) p=0.01, ***) p=0.001

Dansk sammenfatning

Mette Gørtz og Elvira Andersson

Normering i daginstitutioner og pædagogers sygefravær

Dette papir undersøger, om arbejdspres – målt ved de kommunale normeringer – har betydning for sygefraværet på daginstitutionsområdet. Undersøgelsen ser på det langvarige sygefravær – dvs. sygefraværsperioder over 14 dage – og det korte sygefravær, dvs. kortere fraværsperioder, der ikke involverer sygedagpengerefusion. For det lange sygefravær ses på perioden 2002-2006, mens vi for det korte sygefravær kun har oplysninger for 2005-2006. Der opstilles og testes en empirisk model, hvor sygefraværet for den enkelte person ansat i en daginstitution formuleres som en funktion af arbejdspres (normering i kommunen) samt en række andre karakteristika vedr. arbejdspladsen, og der kontrolleres endvidere for personlige karakteristika, som kan tænkes at have indflydelse på sygefraværet.

I papiret diskuteres det indgående, om og under hvilke forudsætninger en eventuel (positiv) korrelation mellem normeringer og sygefravær kan siges at være udtryk for en kausal sammenhæng. Det er tænkeligt, at institutioner/kommuner, der tilbyder gode arbejdsforhold, herunder lave normeringer, får relativt flere ansøgninger og dermed har bedre mulighed for at vælge nye medarbejdere med et stærkt helbred og uden tidligere sygefraværsperioder end institutioner/kommuner med mindre attraktive normeringer. Hvis det er tilfældet, vil en eventuel positiv korrelation mellem normeringer og sygefravær ikke nødvendigvis udelukkende være udtryk for, at gode (lave) normeringer gør medarbejderne mindre syge, men kan også være udtryk for, at kommuner med gode normeringer i højere grad har haft mulighed for at ansætte medarbejdere med et grundlæggende godt helbred og en lav sygefraværstilbøjelighed. Det såkaldte selektionsproblem består således i, at raske personer ansættes i de kommuner, hvor arbejdsforholdene er mest tiltrækkende. Selektionsproblemet kan være stort, hvis arbejdsgiverne har relativt gode muligheder eller evner for at aflæse, om en potentiel medarbejder har et godt helbred og en lav tilbøjelighed til at blive sygemeldt. Hvis selektion skønnes at være omfattende, vil en almindelig korrelation mellem normeringer og sygefravær ikke nødvendigvis kunne fortolkes som en årsagssammenhæng. Selektionsproblemet kan håndteres ved hjælp af forskellige statistiske metoder, hvoraf flere er blevet afprøvet i den empiriske undersøgelse.

Sandsynligheden for at blive langtidssyg estimeres på to forskellige måder: ved hjælp af en fixed effects logit estimator, hvor paneldimensionen i data udnyttes, og ved hjælp af en instrumentvariabelestimation. Instrumentvariabelestimationerne indikerer, at for ansatte i børnehaverne er der en positiv og signifikant sammenhæng mellem normeringer og sandsynligheden for at blive langtidssygemeldt i perioden efter 2005. En stigning i normeringen på 1 (et barn pr. ansat) for en gennemsnitlig børnehave er ensbetydende med, at risikoen for at

blive langtidssygemeldt (over 14 dage) stiger med 4 procentpoint (hvilket skal ses i forhold til en gennemsnitlig risiko for langtidssygemelding på 10%).

Længden af langtidssygdom målt i antal dage analyseres ved hjælp af en fixed-effects negative binomial estimator (FE Negbin) og ved hjælp af en lineær model (2SLS), hvor instrumenterne pasningsgaranti og andel af borgerligt-liberale i kommunalbestyrelsen optræder som instrumenter. Vi finder ikke nogen statistisk signifikant sammenhæng mellem normeringer og langtidssygefravær målt i dage.

Endelig analyseres omfanget af det korte sygefravær for perioden 2005-2006. Estimationerne vha. instrumentvariabelmetoden tyder på, at der er en positiv og signifikant sammenhæng mellem normeringer og kort sygefravær for vuggestuer. En stigning i normeringen på 1 indebærer en stigning i det korte sygefravær på 16 timer om året (i forhold til et gennemsnitligt årligt korttidssygefravær på 55 timer).

Resultaterne er blandede, men tyder overvejende på, at der er en positiv sammenhæng mellem normeringer og sygefravær. De besparelser på lønomkostningerne, som kommunerne kan opnå ved at øge normeringerne, skal således sættes op imod ekstra omkostninger til sygedagpenge, vikardækning for sygemeldte, en stigning i antal personer på førtidspension mv. En simpel cost-benefit-analyse i papiret indikerer, at op mod 20% af besparelsen på de offentlige budgetter ved en stigning i normeringerne vil gå til ekstra sygedagpenge, vikardækning og førtidspension. Hertil skal lægges produktionstab for samfundet, når arbejdsudbuddet reduceres som følge af øget tilgang til førtidspension, tabte skatteindtægter som følge heraf, dårligere kvalitet i børnepasningen samt forringet velfærd for de familier, der er brugere af daginstitutionerne. Disse omkostninger er ikke værdisat i cost-benefit-analysen.

Child-to-Teacher Ratio in Day Care and Teacher Sickness Absenteeism

This paper analyses whether work pressure measured by the child-to-teacher ratio, i.e. the number of children per teacher in an institution, affects teacher absenteeism due to sickness in Danish day-care institutions. We control for individual teacher characteristics like for instance education and family background. Furthermore, we investigate the role of other characteristics at the workplace level like for instance the size of the institution, the proportion of the staff who are trained preschool teachers, family background characteristics of the children in the preschool etc. Our estimation results indicate that for preschool teachers, the risk of becoming long-term ill is positively related to the child-to-teacher ratio for 2005-2006. Furthermore, we look at how the extent of short-term absence is related to the child-to-teacher ratio. Our estimation results indicate that nursery-care teachers' sickness absence is positively related to their work pressure.

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