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Incentives and Selection in Cyclical Absenteeism

by

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Abstract

Procyclical absenteeism might be due to higher sick-rates of marginal workers, or a consequence of procyclical sick-report incentives. These hypotheses predict opposite signs for the correlation between sick-rates and shares of temporary contracts. This is the case, when the share of temporary contracts is a proxy for the share of marginal workers, and an indicator of stronger incentives for job presence of temporary employees who have generally weaker job security than those on permanent contracts. Using industry-region panel data, we find a stable negative correlation between sick-rates and shares of temporary contracts implying that procyclical sick-rate is compatible with the idea that sick-report incentives are procyclical.

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1 Introduction

Several studies report a negative correlation between aggregate sickness absence and the unemployment level.[1] There are two main theoretical explanations for a negative relationship between unemployment and sickness absence. First, unemployment might affect sickness absence behavior. If absenteeism increase the risk of job loss, a higher unemployment rate will then reduce the propensity to report sick. A second mechanism is related to the sickness propensity of the marginal workers entering or leaving the working population in various states of the business cycle. When employers can choose whom to lay-off, the most absence-prone workers are more likely to be laid off in an economic downturn (Leigh (1985)). This implies that in downturns, the employees' absence rate falls. Analogously, the absence rate increases in times of falling unemployment when the marginal individual leaving unemployment to employment has an above average sick-absence rate.

Based on Swedish establishment data aggregated for 14 industries in 3 regions, we observe a positive relation between employment and sick absence implying that workers in expanding industries and regions have higher sick-rates. This is in line with previous studies where unemployment and sick-rates are found to be negatively correlated. Given such a relation one needs to discriminate between the hypothesis that this correlation is due to changes in worker (sick-report) behavior or a result of higher sick-rates of marginal workers and thus a worker-composition effect. Though several studies analyze the role of economic incentives for worker sick absenteeism no previous study deals with empirical testing of

[1]See e.g. Leigh (1985) and Kaivanto (1997). For Swedish studies see Lantto (1992) and Bäckman (1998).

the mechanisms behind procyclical sick absenteeism.[2]

The purpose of this paper is to discriminate between the hypotheses outlined above by investigating the relation between the share of temporary contracts and the sick-rate. The labor protection law in Sweden makes it easier to reduce employment by not renewing temporary contracts rather than facing the costs of firing workers with permanent contracts. New employment contracts are often on a temporary basis. Furthermore, we observe that the share of temporary contracts in expanding establishments is twice this share in contracting establishments. Share of temporary employment in an industry-region cell is thus a good indicator of the share of marginal workers. The selection effect (changes in worker-composition) hypothesis would thus imply a positive correlation between sick-rates and the share of temporary contracts.

The worker incentive effect hypothesis would on the contrary imply a negative correlation between these variables since workers on temporary contracts experience weaker job-security and thus would tend to have less sick absence. This is the case when a low sick-rate improves the chances of a renewed contract. These two hypotheses predict opposite signs of the estimated coefficient for temporary contracts. This enables us to discriminate between these two rival hypotheses.

Our results imply that the sick-rate and the share of temporary contracts are negatively and significantly correlated. This correlation is basically unchanged when we take into account the share of females, young workers, employment as well as industry-region effects. We also control for time-specific effects meant to capture

[2]For the role of economic incentives and sick absenteeism see e.g. Barmby *et al.* (1991, 1994, 1995), Johansson and Palme (1996), Winkelman (1999), Brown *et al.*(1999)

all general business-cycle effects and changes in the social insurance system. Our results do not support the selection effect hypothesis, they indicate instead that worker-behavior is influenced by the type of employment contract implying that less job-security is associated with lower sick-rate. Finally, our sensitivity analysis regarding endogeneity of temporary contracts as well as persistence in sick-rates leaves our results basically unchanged.

The remainder of the paper is as follows. Section 2 deals with data description and the empirical set-up. Our estimation results are reported in Section 3. In section 4 we deal with the dynamics of sick-rates and endogeneity of temporary contracts applying dynamic panel data methods. The paper is concluded in Section 5.

2 Data and empirical set up

The data are from the Short Term Employment Statistics[3] collected by Statistics Sweden (*Statistiska Centralbyrån*, SCB). We use quarterly information on employment stocks and absenteeism for a panel of around 10,000 establishments in the non-agricultural private sector, during the period 1989:1-1999:4. All measures of employment are reported separately for men and women and separately for temporary and permanent employment contracts. Information regarding sickness absence refers to the number of employees absent from work due to sickness a given day.

The sample is drawn from the population of private sector establishments of all sizes in Sweden and stratified according to industry affiliation and establishment size. Establishments with 100 employees or more are sampled with probability one. In order to include

[3]Kortperiodisk sysselsättningsstatistik.

newly started establishments and avoid attrition due to establishments exit, 10 percent of the sample is updated every year for the period 1989-1994. After the first quarter of 1995 the sample is updated every six months. The samples are randomly divided into three groups. Each group reports the requested information for only one month in a quarter. For example one third of establishments report employment in a certain day in January while the establishments in the other two groups report employment in a certain day in February and March respectively.

This sampling strategy implies that the establishment level observations refer to different points in time for different establishments and are not perfectly comparable. On the aggregate-level, however, the information refers to the whole quarter since it is based on three measurements in every quarter. Moreover, the panel structure of the data on the establishment level is not satisfactory when considering longer periods. Small establishments enter and leave the data frequently while large establishments are present for most of the time.

To avoid measurement problems and obtain results referring to the total population we aggregated the establishment level data to a panel data set containing 42 industry-region cells (14 industries times 3 regions). An observation is thus an industry-region cell which we follow quarterly during the period 1989-1999 in a balanced panel yielding 1,848 observations in total. The regional aggregation is done with three regions including: big cities, forest counties and all other counties. The industry classification refers to an aggregation that allows comparison of the Swedish industry classification standard from 1969 to that of 1992.[4]

[4]The aggregated data can be obtained from the authors on request. The underlying establishment level data can be obtained from Statistics Sweden, or

Table 1. Sample statistics, panel of 14 industries in 3 regions, 1989:1 - 1999:4. Number of observations: 1,848.

	Mean	SD	Min	Max
Sick-rate	4	2	0.5	17
Employment (Thousands)	47	59	0.5	355
Female Share of Employment	37	16	9	68
Share of Temporary contracts	10	7	0.6	48
Share of workers younger than 35	38	11	13	71
Share of workers younger than 25	13	8	0.0	44
Share of workers older than 51	25	7	9	55
Share of workers older than 56	14	4	4	30

Sample statistics are reported in Table 1. The dependent variable in our analysis is the *sick-rate* defined as the share of industry-region cell employees who are reported sick. The absenteeism measure here does not include any temporary or permanent disability pensions financed within the general sick insurance system.

Our explanatory variables are aggregate employment, the share of young workers, females and employees with temporary contracts in each industry-region cell. The data on the share of young workers is computed for 1989 -1998 from the individual data set LINDA (Longitudinal Income Data, Statistics Sweden).[5] These data are then matched on industry-region level. The female workers' share of employment is intended to control for variation in sick-rates between women and men and the share of young is meant to capture variation in sick-rates related to age.

We estimate a standard error-component model on our panel of industry-region cells using our variables as described above. Our

from the authors for replication.

[5]The latest wave of LINDA is from 1998 and therefore we set the age shares for 1999 equal to those of 1998.

basic specification is: $y_{it} = \mathbf{x}_{it}^{'}\beta + u_{it}$; $u_{it} = \alpha_i + \lambda_t + \varepsilon_{it}$, where y_{it} is the sick-rate for industry-region cell i in the time t, $\mathbf{x}_{it}$ is a vector of our basic explanatory variables: employment, share of female employees, share of temporary contracts and share of young workers aged 35 years or younger. The error term is decomposed into an industry-region effect, a time effect and a normally distributed error term with a zero mean.

Measuring the number of sick reported individuals in a certain day capture long term absenteeism with a higher probability compared to short term absenteeism. This should not be a major problem since sick-absence for employees with part or full-time temporary disability pensions are not included in our measure. However, given that employees on temporary contracts are less likely to be on long term sick-absence, long term absenteeism is lower in industries with higher shares of temporary contracts. These types of systematic time-invariant industry-region differences are captured by our industry-region fixed effects.[6] Moreover, we estimate a model including the number of permanent jobs as well as an interaction between the share of temporary contracts and the number of permanent jobs. This estimation allows us to control for differential variation over time in the share of temporary contracts that stems from variation in permanent jobs and its impacts on long term sickness captured by our measure of the sick-rate.

Finally, two problems might lead to inconsistent estimates. First, the share of temporary contracts might be endogenous. This is the case when increased sick-rate leads to increased temporary employ-

[6]Notice that the standard variables determining individual absenteeism such as marital status as well as the number and age of children might differ across industries. This is captured by our fixed effects specification. Our objective is not to explain level differences in absenteeism across individuals or industries, but rather the pattern of changes over time.

ment meant to replace absent employees. Such a positive correlation would make it difficult for us to observe a negative effect of temporary contracts on sick-rates. Second, the dynamic of the sick-rate might be characterized by persistence over time. This is the case when workers' sick report behavior adjusts slowly to changes in incentives and/or changes in the share of marginal workers.

We deal with the endogeneity of the share of temporary contracts and the persistence of sick-rate by applying the techniques suggested by Arellano and Bond (1991) as well as Blundell *et al.* (1998, 2000). This yields GMM-estimates where share of temporary contracts are instrumented to account for endogeneity and the lagged sick-rate is included as explanatory variable to incorporate possible persistence of sick-rate.

3 Results

Results reported in Table 2 indicate that sick-rate is positively correlated to employment in line with previous results indicating a negative correlation between unemployment and sick-rate. Our main concern here is the impact of temporary employment on sick-rates which turns out to be negative and highly significant.

These results seem to be robust controlling for a number of potential determinants of sick-rate. The within estimates presented in Table 2 represent the impact of explanatory variables after controlling for industry-region fixed effects such as industry specific working conditions and the local labor market situation.

Workers on temporary contracts have limitations on duration of sick period due to the duration of their contract. Therefore we expect that the share of long-term sick absenteeism is negatively correlated to share of temporary contracts. As mentioned above, our measure of sick-rate captures long-term sickness with a higher probability compared to short-term sick periods. To deal with this problem we controlled for the number of permanent contracts and the interaction between this variable and the share of temporary contracts. These two variables capture variations in sick-rate that stems from variations in the number of permanent contracts as well as variations in the share of temporary contracts that is induced by variations in permanent contracts. Adding these controls leaves our estimate of the share of temporary contracts qualitatively unaffected indicating that the estimate for the share of temporary contracts represents own-effects (Model 4 in Table 2).

We also explicitly take into account the impact of female rate of employment and share of young workers. The coefficient for the share of female workers turns out to be positive and significant as expected. The coefficient for young workers is highly insignificant in Model 2, Table 2, reflecting the fact that within industry-region variation in the share young workers over time is mainly due to the general demographic trend of aging population captured by our time effects. This is confirmed by the results of estimating a model not including time effects where the coefficient for the share of young is positive and highly significant.

We also experimented with two other measures of age. Running Model 2 with the share of workers older than 50 or 55 yield insignificant estimates for these variables. Notice that the share of temporary contracts and older workers are negatively correlated in cross-section. However, our within estimates use the time variation of the variables. The insignificant estimates for the age variables simply reflect the poor over-time variation in age variables other than those associated with the trend of an aging working population.

We consider time fixed effects by allowing time specific intercepts corresponding to each industry-region cell. A possible time effect is the effect of changes in work-pace producing procyclical sick-rates. Time fixed effects capture all global cyclical effects as well as all symmetric effects of changes in the social insurance system.

Table 2. Sick-rates and temporary employment. employment weighted within estimates.

	Model 1	Model 2	Model 3	Model 4
% Temp. contr.	-0.111**	-0.061**	-0.147**	-0.083**
	(0.012)	(0.009)	(0.009)	(0.011)
Employment*10^{-3}	0.002*	0.005**	-	-
	(0.001)	(0.001)		
% Fem. Employment	0.226**	0.049**	-	-
	(0.015)	(0.011)		
% Age $<$ 35	0.305**	-0.011	-	-
	(0.011)	(0.010)		
Perm. contr. * 10^{-3}	-	-	-	0.019**
				(0.0014)
Perm. contr.*	-	-	-	-0.104**
% Temp. contr. * 10^{-5}				(0.011)
Time fixed effects	-	YES	YES	YES
NxT	1848	1848	1848	1848
R^2 within	0.51	0.79	0.58	0.59

Notes: All models include quarterly dummies to control for seasonal variation. Hausman tests indicate that for all model specifications, the individual effect is correlated with the regressors implying that the fixed-effects specification should be used. Significance at the 5 and 1 per cent level are denoted by * and **.

It is interesting to note that the time fixed effects are powerful in explaining the variation in sick-rate. The within R^2 rises from 0.51

to 0.79. This is not surprising since the health insurance system in Sweden is guided by the principle of having general insurance with very little differentiation, a principle that also affects all changes in the system. In this way, the effects of institutional changes on sick-rate is to a large extent symmetric for individuals in various industries. Moreover, the time effects also capture the business cycle effects which seems to have symmetric effects on sick-rates.

These controls do not alter the coefficient for temporary contracts in any significant way. The estimated coefficient is negative and strongly significant through all variations of our specification. This result rejects the idea that the positive correlation between employment and sick-rate is due to a composition effect when share of temporary contracts works as a proxy for marginal workers entering or leaving employment. Incidence of temporary contract is associated with lower sick-rate, implying that workers' absenteeism behavior might be influenced by whether they are on temporary or permanent contracts.

4 Dynamics of sick-rate and endogeneity of temporary contracts

This section is devoted to a sensitivity analysis examining the robustness of our results reported in the previous section. The sick-report behavior of workers might be persistent over time implying that we should include a lag dependent variable in our specification and estimate a dynamic model. We use a GMM estimator to estimate our dynamic model instrumenting the lagged dependent variable. This estimator is denoted as $GMM - IV_y$. Furthermore we account for endogeneity of the share of temporary contracts by instrumenting this variable with its lagged values. Instrumenting both the lagged dependent variable and the share of temporary contracts yields our $GMM - IV_{y,x}$ estimates.[7]

For simplicity, in this section we use annual data. This is motivated by the fact that obtaining GMM estimates with large number of time periods is very demanding, due to the instrumenting procedure involved. This should not have any major impacts on our results, since our within estimates based on annual and quarterly data are qualitatively the same. However, to examine the role of possible biases in our within estimates we compare the GMM estimates with within and OLS estimates on annual data.

We re-examine the own-effect of temporary contracts, i.e. the estimate of the temporary contracts when including the number of permanent contracts and the interaction between the share of temporary contracts and the number of permanent contracts as done in Model 4 in Table 2. Using our various GMM-estimators, we find a similar pattern as in the case of the within estimator presented in Table 2.

The estimated coefficient for the lagged dependent variable indicate substantial persistency in the rate of sick absenteeism. Test statistics for AR (1) and AR (2) indicate that the residuals are not random walk and imply absence of serial correlation. According to the Sargan test we can not reject that the instruments are uncorrelated with the errors implying validity of our instruments.

The estimated coefficients for employment seem to be rather unstable. This is natural due to the high correlation of lagged and present sick-rates, implying that the correlation between employment and present sick-rate is partly captured by the lagged sick-rate. For our purpose, the main point is that the effect of tem-

[7]See Arellano and Bond (1991) and Blundell and Bond (1998, 2000)

porary contracts is always negative and significant across various model specifications.

Table 3. Sick-rates and Temporary Employment using annual data. Robust standard errors in parentheses.

	OLS	Within	GMM			
			IV_y	IV_y	$IV_{y,x}$	$IV_{y,x}$
			DIF	SYS	DIF	SYS
Sick-rate (-1)	0.810**	0.554**	0.725**	0.706**	0.704**	0.691**
	(0.018)	(0.039)	(0.087)	(0.124)	(0.087)	(0.082)
% Temp.	-0.025**	-0.056**	-0.101**	-0.055*	-0.090*	-0.127**
	(0.008)	(0.019)	(0.038)	(0.025)	(0.036)	(0.045)
Empl. $*10^{-3}$	-0.001	0.000	0.002	-0.002*	0.003	-0.004*
	(0.001)	(0.002)	(0.003)	(-0.001)	(0.003)	(0.002)
% Fem. Empl.	0.002	0.017	0.049**	0.007	0.047*	0.012*
	(0.002)	(0.015)	(0.022)	(0.004)	(0.024)	(0.005)
% Age < 35	0.014**	0.034**	-0.000	0.021*	-0.002	0.052**
	(0.005)	(0.013)	(0.020)	(0.011)	(0.021)	0.018
NxT	378	378	378	420	378	420
AR(1)	-	-	-2.77**	-2.64**	-2.77**	-2.90**
AR(2)	-	-	-0.50	-0.29	-0.48	-0.43
Sargan Test	-	-	33	31	33	30

NOTES: Significance at the 5 and 1 per cent level are denoted by * and **. All models include yearly dummies. The $GMM - IV_y$ estimator uses instruments for the lagged dependent variable. The $GMM - IV_{y,x}$ estimator uses instruments for the lagged dependent variable as well as instruments for the share of temporary contracts. DIF denotes the GMM-estimator using the Arellano & Bond (1991) procedure where first-differences are instrumented using levels. SYS refers to the GMM-estimator using the Blundell & Bond (1998, 2000) procedure where first-differences and levels are used as instruments. GMM-estimates are obtained using DPD for Ox (see Doornik (1999)).

The various estimations presented in Table 3 always yield a negative and significant coefficient for the share of temporary contracts. The size of the coefficients range between -0.12 and -0.05. This is however a result of different estimators properties as well as the instrumenting procedure involved. For our purpose, the essential matter is that the coefficient is always negative and highly significant regardless of which estimator we use. Results reported in Table 3 unambiguously confirm the pattern that the share of temporary contracts and sick-rate is negatively correlated.

5 Final remarks

We tested the hypothesis whether the observed procyclical pattern of sick-rates is due to changes in workers sick-rate incentives as the conditions in the labor market change or if this simply reflects changes in worker-composition when marginal workers with higher sick absenteeism enter or leave the working population.

The share of temporary contracts being a proxy of marginal workers would then be positively correlated to sick-rates according to the latter hypothesis. On the other hand, employees on temporary contracts have stronger incentives for presence at job when this affects their future employment chances. The share of temporary workers should then be negatively correlated with sick-rate, when workers sick-report incentives affect the aggregate sick-rate.

Our results indicate a negative relation between the share of temporary contracts and the sick-rate in industry-region cells during the nineties in Sweden. Both the selection and incentive mechanism might be present behind the pattern of sick-rate over time, but our results are in line with the idea that incentive-effects dominate. The evidence seem to support the explanation that procyclical sick-rate is due to changes in worker sick-rates report incentives. The correlation between the share of temporary contracts and sick-rates is robust when controlling for industry-region effects, global time effects, share of females, age structure as well as variation in permanent jobs and changes in the share of temporary contracts resulted from variations in permanent jobs.

We also estimated a dynamic model taking into account possible persistence in the sick-rate and instrumented the share of temporary contracts by its lagged values to account for the possibility of endogeneity bias. All results confirm a negative correlation between sick-rate and the share of temporary contracts indicating that workers on temporary contracts have weaker sick-report incentives.

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