Employment status and perceived health condition: longitudinal data from Italy

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Employment status and perceived health condition: longitudinal data from Italy

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Abstract

Background The considerable increase of non-standard labor contracts, unemployment and inactivity rates raises the question of whether job insecurity and the lack of job opportunities affect physical and mental well-being differently from being employed with an open-ended contract. In this paper we offer evidence on the relationship between Self Reported Health Status (SRHS) and the employment status in Italy using the Survey on Household Income and Wealth; another aim is to investigate whether these potential inequalities have changed with the recent economic downturn (time period 2006-2010).
Methods We estimate an ordered logit model with SRHS as response variable based on a fixed-effects approach which has certain advantages with respect to the random-effects formulation and has not been applied before with SRHS data. The fixed-effects nature of the model also allows us to solve the problems of incidental parameters and non-random selection of individuals into different labor market categories.
Results We find that temporary workers, unemployed and inactive individuals are worse off than permanent employees, especially males, young workers, and those living in the center and south of Italy.
Conclusion Health inequalities between unemployed/inactive and permanent workers widen over time for males and young workers, and arise in the north of the country as well.

Keywords: self-reported health status, employment status, economic crisis, fixed-effects ordered logit model

Background

Job insecurity and the lack of work opportunities have characterized labor markets for the better part of the last decade. There has been a considerable increase of non-standard labor contracts, as opposed to permanent employment, alongside the increase in the unemployment and inactivity rates, especially among young people. In Italy, for workers between 15 and 24 years of age, the share of temporary employment on total employment raised from 26.2% in 2000 to 52.9% in 2012, the unemployment rate raised from 31.2% to 35.2%, and the inactivity rate went from 61.6% to 71.3% OECD (2013).

In this framework, one relevant question in applied works has been whether the perceived job insecurity and being out of the labor market affect the individuals well-being differently

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from being employed with a permanent contract. The available empirical evidence suggests that there is a negative relationship between being employed with a temporary contract and the individual health status and several contributions have also found a negative correlation between long-term unemployment, health condition, and mortality risk. Moreover, the intensity of this relationship has been proved to be strongly differentiated by gender and across countries, especially in relation to the institutional settings that characterize the labor markets.

Switching between unemployment and different contract types has been found to have health effects in West-Germany and Spain and the transition form unemployment to a temporary contract has smaller benefits than transiting into permanent employment Gash et al. (2007). In the British case, both contractual and working conditions affect the workers well-being with marked differences across genders Robone et al. (2011). In addition, the differences in workers self-reported well-being are associated with the satisfaction with job security Dawson and Veliziotis (2013). There is also some evidence from Scandinavian countries: form the self-rated health in the Danish workforce questionnaire for the period 1995-2000, job insecurity causes a decline in self-rated health and that this effect is more intense for female workers Rugulies et al. (2008); temporary employment is also related to a decline in self-rated health and psychological distress in a sub-cohort of the Northern Swedish Cohort in 2007 Waenerlund et al. (2011). Moreover, the total amount of accumulated unemployment during the deep Swedish recession of 1992-1996 is related to elevated all-cause mortality for men and women in the following 6 years and the mortality risk increases with the duration of unemployment Garcy and Vågerö (2012). Results from the German Socio-Economic Panel for the years 1991-2008 suggest a negative correlation between unemployment and health condition as it is associated with higher risks of a heart attack or stroke and with the risk of mental illness Schmitz (2011).

Extensive results have also been gathered for extra-European countries. Studies using the Korea Labor and Income Panel Survey suggest that the deterioration in the health status is related to the contract type for both men and women Kim et al. (2008). A similar study is carried for Japan, although some gender differences emerge Nishikitani et al. (2012). Testing the relationship between working hours, change in type of contract and health after the 2008 economic crisis using a U.S. longitudinal community-based sample suggests that changes in the variable of interest have no significant effect on health behaviors Macy et al. (2013). Meta-analyses show that the unemployed among working-age people have an increased risk of death in the U.S., higher for men than for women Roelfs et al. (2011). Unemployment is also associated with an increased mortality risk for those in their early and middle careers, but less for those in their late careers. In the case of Australia, temporary full-time employment is associated only with poor physical health without affecting the workers psychological well-being Keuskamp et al. (2013). The relationship between employment and health has also been analyzed for the case of Brazil Giatti et al. (2010): after adjusting for individual socio-demographic characteristics, behavioral risk factors and health status, they find that the association of unemployment and socioeconomic characteristics of the neighborhoods in which people live is related with poor self-rated health.

Recently, cross-country analyses have also been carried out with the purpose of highlighting differences within Europe. The association between job insecurity and self-rated health has been studied using cross-sectional data from 16 European countries finding that precarious employment is not associated with poor health only in Belgium and Sweden Laszlo et al. (2010). Institution settings and labor market regulations can explain a significant part of cross-country differences Cottini and Lucifora (2010). A study realized in six countries (United States, Netherlands, England, Finland, Italy, and Spain) shows that living in more deprived neighborhoods is related to higher mortality rates (for all causes) independently of individual socioeconomic characteristics (education and occupation at the individual level), but the relation is not modified by the country context van Lenthe et al. (2005). It has also been investigated how economic changes have affected mortality rates over the past three decades between 1970 and 2007 across differences in government health expenditures for 26 E.U. countries Stuckler et al. (2009). However, there is no consistent evidence across the E.U. that mortality rates increased with unemployment. Long–term unemployment is associated with a greater incidence of suicide: in particular, the risk is greatest in the first five years, and persists at a lower level up to 16 years after unemployment Milner et al. (2013). Moreover, welfare regimes may be important determinants of the employment-health relation and they can result in different consequences for the health effects of insecure and precarious employment: precarious workers in Scandinavian countries report better or equal health status when compared with their permanent counterparts. On the other hand, precarious work in Bismarckian, Southern European, Anglo-Saxon, Eastern European, and East Asian countries is found to be associated with adverse health outcomes, including poor self-rated health, musculoskeletal disorders, injuries, and mental health problems Kim et al. (2012).

Some empirical works are instead specifically focused on the effect of temporary employment on mental health. For the U.S., job insecurity has been found to increase by 50% the level of depressive symptoms Quesnel-Vallee et al. (2010). The relationship between job precariousness and poor mental health has also been analyzed using data from the Psycho-social Factors Survey carried out in 2004-2005 in Spain and employing the response variables form the *Employment Precariousness Scale* (EPRES) Vives et al. (2013). This variable comprises many dimensions of temporary employment (duration, economics deprivation, limited rights, vulnerability and defenselessness in the work place [1]). They find that the association between job insecurity and poor mental health is significant and stronger for women. Data on Slovak and Dutch cities provide information on the association between mental health problems and local unemployment: the interaction is strong in the Netherlands, but absent in Slovakia where citizens from the most favorable neighborhoods have a nearly double the risk of mental health problems than their Dutch counterparts Behanova et al. (2013).

Although these empirical works have assessed the extent of this relationship for many OECD countries, the evidence for the case of Italy is indeed scarce. Only one contribution provides some empirical evidence for the Italian case Carrieri et al. (2012): the study employs the "Multiscopo" (Multiscope) survey issued by the Italian Institute of Statistic matched, by a simulation procedure, with some information (non–health related) on income present in SHIW. Youth temporary employment is analyzed for the years 2004/2005. It emerges that there is a negative

[1] See Vives et al. (2010) for further details

relationship between psychological well-being, happiness and temporary employment especially for young male workers. However, their analysis is referred only to the pre-crisis period.

In this paper, we offer empirical evidence on the relationship between the self-assessed health status and the labor market position in Italy for the first time. In particular, we compare the health conditions of temporary workers, unemployed and inactive individuals with those of workers with fixed income (permanent employees and self-employed workers).

1 Methods

Data source

We use the question on Self Reported Health Status (SRHS) in the panel Survey on Household Income and Wealth (SHIW) issued by the Bank of Italy. For the longitudinal study SHIW about 2,000 households are interviewed every two years. To date, data are collected up until 2010 which is the last issue available. The SHIW includes the question on SRHS only since 2006, so that the analysis of the respondents' self-assessed health condition can be carried out for the period 2006-2010 with 3 time points. For this time window we build an unbalanced longitudinal dataset of $59,294$ observations for which SRHS is observed at least twice. We further restrict the sample to individuals between 15 and 64 years of age and, after dropping outliers and observations with missing values in the variables of interest, we end up with a sample of $37,477$ observations.

SRHS, labeled `SALUT` in the questionnaire, is an ordered variable that takes values between 1 and 5 for increasing health status: it takes value 1 if the respondent answered his/her health is *very poor* in the year of the interview, it takes value 2 for *poor*, 3 for *fair*, 4 for *good* and 5 for *excellent*. The first row of Table 1 contains the sample frequencies for each category of SRHS and shows that the majority of the respondents declare they are in good or excellent health.

[Table 1 about here.]

The employment status is a categorical variable that identifies four possible conditions: we label *permanent* dependent workers with an open–ended contract and self-employed individuals with a stable income[2]; the *temporary* category comprises job contracts such as apprenticeships, on-project jobs, and seasonal jobs [3]; the *unemployment* category includes both unemployed individuals and first–job seekers; finally *inactive* embeds home makers, retired workers[4], and students. The distribution of responses to SRHS by employment status does not exhibit major differences (see the top of Table 1). Nevertheless, the share of inactive individuals choosing the *excellent* category is lower than for the other three classes as well as the share of responses to *good*, decreasing from *permanent* to *inactive*. Moreover, the share for responses *fair*, *poor*, and *very poor* is also increasing from *permanent* to *inactive*.

[2] The information on the working status is provided by the variable `APQUAL` in the SHIW questionnaire.

[3] For self-employment, there is a specific category of `APQUAL` that isolates non-standard contracts. For dependent workers, we cross-reference the categories in `APQUAL` with the variable `CONTRATT`, that takes values 2 and 3 if the job is not permanent.

[4] Individuals who retired because of a disability have, however, been excluded.

Table 1 also shows the respondents' distribution by year, gender, age, and area of residence. The shape of the distribution of responses to SRHS has not dramatically changed with the occurrence of the economic crisis. Furthermore, Table 1 shows that there is a higher concentration of Good responses for women compared to the higher share of Excellent responses for male individuals. Finally, there is a clear difference in the distribution of SRHS between the North and the Rest of Italy: while in the North the majority of responses are equally shared between the Good and the Excellent category, for the rest of the country there is a tendency to concentrate responses on the Good category.

We also present descriptive statistics for the household net income[5], the household wealth and the regional unemployment rate[6]. Although they are not of primary interest in our analysis, income and wealth will be included (in thousands of euros) in the model specification as proxy for economic deprivation, while the unemployment rate will give a measure of relative deprivation. For readability of the descriptive statistics, we categorized income, wealth, and regional unemployment rate by their quartiles: as expected, the relative distribution of the responses shifts towards higher categories of SRHS as income and wealth increase, while it progressively concentrates on the Good category (leaving Excellent) for increasing values of the local unemployment rate.

Statistical analysis

The nature of SRHS is such that it needs to be modeled as an ordinal response variable and therefore non–linear models must be employed for estimation. In applied works, it is customary to consider the ordinal response variable as the results of a categorization mechanism that limits the observability of a latent continuous variable, as the health status can be. Therefore, we set up our model specification by defining the unobservable perceived health status as the latent continuous variable y^*_{it} for individual i at time t that is based on a linear combination of individual covariates collected in the column vector $\boldsymbol{x}_{it}$ and unobservable characteristics represented by the random variables α_i and ε_{it}:

$$y^*_{it} = \boldsymbol{x}'_{it}\boldsymbol{\beta} + \alpha_i + \varepsilon_{it}, \quad \text{for} \quad i = 1, \dots, N \ t = 1, \dots, T. \tag{1}$$

In the above expression, the unobservable individual effect α_i is allowed to be correlated with $\boldsymbol{x}_{it}$. The observable ordinal health status is related to the latent variable y^*_{it} by the following observational rule:

$$y_{ik} = k \quad \text{if} \quad \tau_k < y^*_{it} \leq \tau_{k+1}, \ k = 1, \dots, K, \tag{2}$$

where the threshold parameters τ_k, $k = 1, \dots, K$, are strictly increasing with $\tau_1 = -\infty$, $\tau_{K+1} = \infty$, and K is the number of response categories.

With cross-section data, the parameters $\boldsymbol{\beta}$ in (1) and the threshold parameters τ_k in (2) are estimated by Maximum Likelihood using the common ordered logit/probit models provided that the effects α_i are ruled out. However, the longitudinal structure of our dataset allows us

[5] Since it is computed using both dependent labor income and self-employment, income can be negative.

[6] We matched data on the regional unemployment rate published by the Italian Institute of Statistics with SHIW data at the regional level.

to properly take into account the presence of the time-invariant individual unobserved heterogeneity effects α_i. With longitudinal data, one can either estimate a random-effects ordered probit/logit model or a fixed-effects ordered logit model. Choosing one or the other implies different assumptions on the effects α_i. In order to consistently estimate a random-effects model, α_i needs to be independent of $\boldsymbol{x}_{it}$ and assumptions on the joint distribution of α_i and ε_{it} must be made; in contrast, the fixed-effects ordered logit model does not require these assumptions as the individual time-invariant unobserved effects cancel out with a suitable transformation that is illustrated in the following. The fixed-effects model also provides an estimator robust to misspecification of the distribution of α_i. Note that the estimator is less efficient that the one obtained with the random-effects model when the distributional assumptions on α_i are correct.

A fixed-effects ordered logit model is a better choice for our applications as it solves the problem of other possible sources of bias due to unobserved heterogeneity. Most importantly, there is the issue of self-selection into employment conditions De Cuyper et al. (2008); Carrieri et al. (2012): unobserved characteristics, possibly also health related, may non-randomly group workers into different contract types, unemployment and inactivity, and gives rise to a bias in their effect on the health status. By assuming that such heterogeneity is time-invariant in the time span considered, the bias is eliminated if an ordered fixed-effects logit model is used for the estimation.

The estimation of an ordered fixed-effects logit model can be reduced to the estimation of a fixed-effects binary logit model Andersen (1970, 1972); Chamberlain (1980) once the J categorical responses have been collapsed into two categories. A consistent estimator can be obtained by conditioning each likelihood contribution on a sufficient statistic which is the sum of the individual outcomes over time. The parameters, associated with time-varying covariates, can then be estimated by Conditional Maximum Likelihood (CML).

By assuming that the the error terms ε_{it} are IID standard logistically distributed conditionally on $\boldsymbol{x}_{it}$ and α_i, the probability that y_{it} takes value k for individual i at time t is

$$Pr(y_{it} = k|\boldsymbol{x}_{it}, \alpha_i) = \Lambda(\tau_{k+1} - \boldsymbol{x}'_{it}\boldsymbol{\beta} - \alpha_i) - \Lambda(\tau_k - \boldsymbol{x}'_{it}\boldsymbol{\beta} - \alpha_i), \tag{3}$$

where Λ is the standard logistic cumulative distribution function. Following standard notation Baetschmann et al. (2011), d_{it}^k is the binary dependent variable defined as $d_{it}^k = \mathrm{I}(y_{it} > k)$, that is the dichotomization of y_{it} at the cutoff k. It follows that $P(d_{it} = 0) = \Lambda(\tau_{k+1} - \boldsymbol{x}'_{it}\boldsymbol{\beta} - \alpha_i)$ and $P(d_{it} = 1) = 1 - \Lambda(\tau_{k+1} - \boldsymbol{x}'_{it}\boldsymbol{\beta} - \alpha_i)$. The sum over all elements of $\boldsymbol{d}_i^k = \left(d_{i1}^k, \ldots, d_{iT}^k\right)$ is a sufficient statistic for α_i, as in the binary model, and the thresholds τ_k: by conditioning on $\sum_{t=1}^T d_{it}^k = d_{i+}^k$ it can be shown that the α_i can be eliminated[7].

The joint density of $\boldsymbol{d}_i^k$ is

$$f_{ki}\left(\boldsymbol{d}_i^k \middle| \sum_{t=1}^T d_{it}^k = d_{i+}^k\right) = \frac{\exp\left(\sum_{t=1}^T d_{it}^k \boldsymbol{x}'_{it}\boldsymbol{\beta}\right)}{\sum_{\boldsymbol{b}\in B_{d_{i+}^k}} \exp\left(\sum_t b_{it}\boldsymbol{x}'_{it}\boldsymbol{\beta}\right)}$$

where B_c is the set of all possible sequences of 0s and 1s for which the sum of T binary outcomes is equal to $\sum_{t=1}^T d_{it}^k = d_{i+}^k$, $B_{d_{i+}^k} = \{\boldsymbol{b}_i : \sum_{t=1}^T b_{it} = \sum_{t=1}^T d_{it}^k = d_{i+}^k\}$. The sample conditional

[7] See Chamberlain (1980) and Baetschmann et al. (2011) for further details.

log-likelihood is

$$\ell(\boldsymbol{\beta}) = \sum_{i=1}^{N} \sum_{k=2}^{K} \log f_{ki}(\boldsymbol{\beta}). \tag{4}$$

The vector $\widehat{\boldsymbol{\beta}}$ resulting form the maximization of (4) is a consistent estimator of the parameters in (1). Only individuals with $d_{i+}^{k} < T$ and $d_{i+}^{k} > 0$ contribute to the log-likelihood. In addition, the fixed-effect nature of the model is such that parameters associated with time-invariant covariates cannot be estimated because they are not identified.

Finally, we test the fixed-effects type against the random-effects specification of the model by means of the Hausman test Hausman (1978): under the null hypothesis of correct specification of the joint distribution of α_i and ε_{it}, both the fixed-effects and the random-effects estimators are consistent but the latter is more efficient; under the alternative, only the fixed-effects estimator is consistent. However, the proposed test is not valid with heteroskedasticity and when time dummies are included in the model specification. Therefore, we estimate with a random-effects[8] ordered logit model the auxiliary regression

$$y_{it}^{*} = \boldsymbol{x}_{it}'\boldsymbol{\beta} + \bar{\mathbf{z}}_i'\xi + \alpha_i + \varepsilon_{it}, \quad \text{for} \quad i = 1, \ldots, N \;\; t = 1, \ldots, T$$

where $\boldsymbol{x}_{it}$ contains time-varying and time-invariant covariates and $\bar{\mathbf{z}}_i$ are the time averages of all the time-varying regressors Wooldridge (2010). The Hausman test statistic is then computed as a Wald test for $H_0 : \xi = \mathbf{0}$ using the panel robust covariance matrix estimator.

2 Results and discussion

Table 2 reports the estimation results of two models for the SRHS variable estimated by using the Fixed-Effects (FE) and the Random-Effects (RE) ordered logit models. The first model, M1, is estimated by using a specification that includes dummies for the employment condition and time dummies separately, while the specification of second model, M2, is augmented by their interaction terms. We include the household annual income and wealth in order to control for the effect of economic deprivation and we add their quadratic terms to capture the well known concave relationship with these covariates. Moreover, the regional unemployment rate serves as a proxy of relative deprivation. Only if the estimation is carried out by the RE ordered logit model, we include the age of the individual in 2006 because it is otherwise not identified.

[Table 2 about here.]

The estimation results for model M1 in Table 2 show that the information produced by the FE and the RE specifications is rather coherent: being employed with a temporary contract, unemployed or inactive in the labor market has a significant negative effect on the perceived health condition compared to be being a permanent worker. Moreover, the coefficients related to the time dummies are also negative and statistically significant suggesting that the perceived health status has possibly decayed during the recent economic downturn. The only difference in

[8] See Greene and Hensher (2010) for a detailed illustration of the estimation procedure.

sign between the FE and RE models concerns the regional unemployment rate. The empirical evidence on the effect of the labor market context, such as local unemployment, on health is rather ambiguous. Our FE model shows a positive relationship between local unemployment and self-reported health, a result that has been found for some European countries as well Strandh et al. (2011): being a temporary worker, unemployed or not employed is less stigmatizing in contexts where unemployment is common. Finally, the p-value of the Hausman test leads us to reject the null hypothesis of correct specification of the RE assumptions, meaning that the results obtained by FE estimation are more reliable.

The columns labeled model M2 in Table 2 refer to the estimation of ordered logit models where the latent health status is specified as

$$y_{it}^{*} = \sum_{j} x'_{itj}\beta_j + \gamma_t D_t + \sum_{j} D_t x_{itj}\phi_{tj} + \alpha_i + \varepsilon_{it}, \quad i = 1, \ldots, N,$$

where $j = \{\text{Temporary, Unemployed, Inactive}\}$, $t = 2008, 2010$ and the control variables have been omitted for brevity. This specification allows us to investigate whether the effect of the employment status on health has changed over time, possibly giving an indication of whether inequalities have strengthened with the occurrence of the economic crisis. However, the coefficients associated with the interaction terms ϕ_{tj} are not directly interpretable and further diagnostics are needed. In particular, we want to test whether the effect of being in a certain labor market condition in 2006, β_j, is the same in 2008 or 2010. To this aim, we perform a Wald test for the null hypothesis $H_0 : \beta_j - \gamma_t - \phi_{tj} = \delta_{jt} = 0$ and the value of test statistic is then compared with a χ_1^2 distribution. Results are displayed in Table 3: the inequality in the health status has widened between all the categories considered and permanent workers during the economic crisis. In particular, inequalities start to grow since 2008 for temporary and inactive workers, while they widen starting in 2010 for the unemployed.

[Table 3 about here.]

As discussed above, there is extensive empirical evidence on differences in the relationship between the employment status and health between female and male workers. Gender inequalities in health are usually tied to the fact that females have a higher life expectancy than males. Moreover, our data suggest that the distribution of responses to SRHS by employment status is somewhat more uniform for women than for men[9]. Therefore, we estimate two separate models for men and women in order to detect such inequalities in the health status. Table 4 reports the estimation results obtained with the FE ordered logit model for the two sub-samples.

[Table 4 about here.]

Male workers with a temporary contract, unemployed or inactive in the labor market present a significant decay in the health condition compared to permanent workers; the perceived health

[9]About 55% of responses in sample of female is *good* (the modal value) in any employment category (except for the unemployed with 53%) while the percentage of responses to the *good* category in the male sample (the modal value) is decreasing with the employment condition: 55% of *permanent*, 53% of *temporary*, 52% of *unemployed*, and 46% of *inactive*.

status worsens after 2006 for all individuals and these inequalities significantly increase after 2006 (see Table 5). In contrast, the health status of female workers does not seem to depend on the employment condition neither in the reference year nor afterwards.

[Table 5 about here.]

Since our sample comprises individuals between 15 and 64 years of age, we also divide our sample in three age groups in order to investigate whether inequalities in the self-reported health differ across ages. Table 6 shows the estimation results of the ordered fixed-effects logit models: non–working individuals are worse off than permanent workers in all the groups considered, while only temporary workers younger than 55 exhibit lower well-being than permanent workers.

[Table 6 about here.]

Moreover, Table 7 shows that inequalities in the health condition grow over time only for the 15-34 group for all the employment categories considered, while only inactive individuals between 35 and 54 experience, in 2010, a further decay in the health condition compared to permanent workers.

[Table 7 about here.]

Finally, it is well known that the geographical location in Italy is tied to strong socio-economic inequalities, especially concerning labor market protection and opportunities. Moreover, since health care expenditure is in hands of the regional governments, differences in health care emerge across the country, where the north benefits from a more efficient health system compared to the rest of Italy Francese and Romanelli (2011). Therefore, we finally investigate whether there is a territorial heterogeneity in the effect of the employment condition on the perceived health status. Descriptive statistics already suggest that there is a strong dichotomy between the north and the rest of Italy in terms of self-assessed health conditions (see Table 1).

Table 8 reports the estimation results of the FE models by area of residence, that is the north of Italy and the rest of the territory. We have to consider the center and south of Italy together because of quasi-collinearity issues; however Table 1 shows that their health patterns are indeed similar. We find evidence that there is a negative effect of unemployment on health in the north of Italy, while temporary, unemployed and inactive workers all exhibit a significant lower health condition compared to permanent workers in the south of Italy.

[Table 8 about here.]

For temporary workers in the north of Italy, health inequalities do not arise over time while they grow for the unemployed and appear for inactive individuals (see Table 9). In the rest of Italy, differences in the perceived health condition widen over time for temporary, unemployed and inactive individuals.

[Table 9 about here.]

3 Conclusions

The incidence of non-standard labor contracts on permanent employment has been steadily increasing and job opportunities for the unemployed and first–job seekers have drastically diminished in the last decade. In this framework, work arrangements and job deprivation have recently become the focus of many empirical contributions aimed at assessing whether an increasing degree of job insecurity and lack of job opportunities are related to worse physical and mental health. This question is particularly relevant in the context of a dual labor market where on one side employees with open–ended contracts enjoy the benefits of a high degree of protection, while on the other temporary workers are exposed to high job insecurity and a very low, if none, degree of employment protection. Moreover, welfare is not always able to adequately sustain workers during spells of unemployment, whose health condition may deteriorate due to the economic deprivation.

We find that, as for the majority of other case studies in this literature, there is a negative relationship between job insecurity/unemployment and health: there is a significant negative effect on the perceived health condition of being a temporary worker, or not working at all, compared to permanent and self employed workers. On average, the health condition decays and inequalities between temporary workers/non-working individuals and the permanent workers grow over time. The same pattern applies for males and young workers in our sample, while there seem to be no strong differences in perceived health for females in different labor market conditions. The negative dependence between job insecurity, inactivity and self-assessed health is stronger in the center and south of Italy than it is the north. By contrast, such inequalities seem to raise in the whole country with the occurrence of the economic crisis.

List of abbreviations

SRHS: Self Reported Health Status; SHIW: Survey on Household Income and Wealth

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References

Andersen, E. B. (1970). Asymptotic properties of conditional maximum-likelihood estimators. *Journal of the Royal Statistical Society. Series B (Methodological)*, pages 283–301.

Andersen, E. B. (1972). The numerical solution of a set of conditional estimation equations. *Journal of the Royal Statistical Society. Series B (Methodological)*, pages 42–54.

Baetschmann, G., Staub, K., and Winkelmann, R. (2011). Consistent estimation of the fixed effects ordered logit model. IZA Discussion Papers 5443, Institute for the Study of Labor (IZA).

Behanova, M., Nagyova, I., Katreniakova, Z., van Ameijden, E. J., van Dijk, J. P., and Reijneveld, S. A. (2013). The effect of urban-area unemployment on the mental health of citizens differs between slovak and dutch cities. *Health & Place*, 24:210 – 215.

Carrieri, V., Novi, C. D., Jacobs, R., and Robone, S. (2012). Well-being and psychological consequences of temporary contracts: the case of younger italian employees. Working Papers 079cherp, Centre for Health Economics, University of York.

Chamberlain, G. (1980). Analysis of covariance with qualitative data. *Review of Economic Studies*, 47(1):225–238.

Cottini, E. and Lucifora, C. (2010). Mental health and working conditions in european countries. IZA Discussion Papers 4717, Institute for the Study of Labor (IZA).

Dawson, C. and Veliziotis, M. (2013). Temporary employment, job satisfaction and subjective well-being. Working Papers 1309, University of the West of England.

De Cuyper, N., De Jong, J., De Witte, H., Isaksson, K., Rigotti, T., and Schalk, R. (2008). Literature review of theory and research on the psychological impact of temporary employment: Towards a conceptual model. *International Journal of Management Reviews*, 10(1):25–51.

Francese, M. and Romanelli, M. (2011). Healthcare in italy: expenditure determinants and regional differentials. Temi di discussione 828, Bank of Italy.

Garcy, A. M. and Vågerö, D. (2012). The length of unemployment predicts mortality, differently in men and women, and by cause of death: A six year mortality follow-up of the swedish 1992–1996 recession. *Social Science & Medicine*, 74(12):1911–1920.

Gash, V., Mertens, A., and Gordo, L. R. (2007). Are fixed-term jobs bad for your health?: A comparison of west-germany and spain. *European Societies*, 9(3):429–458.

Giatti, L., Barreto, S. M., and César, C. C. (2010). Unemployment and self-rated health: neighborhood influence. *Social Science & Medicine*, 71(4):815–823.

Greene, W. H. and Hensher, D. A. (2010). *Modeling ordered choices: a primer*. Cambridge University Press.

Hausman, J. A. (1978). Specification tests in econometrics. *Econometrica: Journal of the Econometric Society*, pages 1251–1271.

Keuskamp, D., Ziersch, A. M., Baum, F. E., and LaMontagne, A. D. (2013). Precarious employment, psychosocial working conditions, and health: Cross-sectional associations in a population-based sample of working australians. *American Journal of Industrial Medicine*.

Kim, I.-H., Muntaner, C., Vahid Shahidi, F., Vives, A., Vanroelen, C., and Benach, J. (2012). Welfare states, flexible employment, and health: a critical review. *Health policy*, 104(2):99–127.

Kim, M.-H., Kim, C.-y., Park, J.-K., and Kawachi, I. (2008). Is precarious employment damaging to self-rated health? results of propensity score matching methods, using longitudinal data in south korea. *Social Science & Medicine*, 67(12):1982–1994.

Laszlo, K. D., Pikhart, H., Kopp, M. S., Bobak, M., Pajak, A., Malyutina, S., Salavecz, G., and Marmot, M. (2010). Job insecurity and health: A study of 16 european countries. *Social Science & Medicine*, 70(6):867 – 874.

Macy, J. T., Chassin, L., and Presson, C. C. (2013). Predictors of health behaviors after the economic downturn: A longitudinal study. *Social Science & Medicine*, 89(0):8 – 15.

Milner, A., Page, A., and LaMontagne, A. D. (2013). Long-term unemployment and suicide: a systematic review and meta-analysis. *PloS one*, 8(1):e51333.

Nishikitani, M., Tsurugano, S., Inoue, M., and Yano, E. (2012). Effect of unequal employment status on workers health: Results from a japanese national survey. *Social Science & Medicine*, 75(3):439–451.

OECD (2013). Stat extracts. OECD Stat Extracts. Available at http://www.oecd.org/statistics/.

Quesnel-Vallee, A., DeHaney, S., and Ciampi, A. (2010). Temporary work and depressive symptoms: A propensity score analysis. *Social Science & Medicine*, 70(12):1982 – 1987.

Robone, S., Jones, A., and Rice, N. (2011). Contractual conditions, working conditions and their impact on health and well-being. *The European Journal of Health Economics*, 12(5):429–444.

Roelfs, D. J., Shor, E., Davidson, K. W., and Schwartz, J. E. (2011). Losing life and livelihood: a systematic review and meta-analysis of unemployment and all-cause mortality. *Social Science & Medicine*, 72(6):840–854.

Rugulies, R., Aust, B., Burr, H., and Bültmann, U. (2008). Job insecurity, chances on the labour market and decline in self-rated health in a representative sample of the danish workforce. *Journal of Epidemiology and Community Health*, 62(3):245–250.

Schmitz, H. (2011). Why are the unemployed in worse health? the causal effect of unemployment on health. *Labour Economics*, 18(1):71–78.

Strandh, M., Novo, M., and Hammarström, A. (2011). Mental health among the unemployed and the unemployment rate in the municipality. *The European Journal of Public Health*, 21(6):799–805.

Stuckler, D., Basu, S., Suhrcke, M., Coutts, A., and McKee, M. (2009). The public health effect of economic crises and alternative policy responses in europe: an empirical analysis. *The Lancet*, 374(9686):315–323.

van Lenthe, F. J., Borrell, L., Costa, G., Roux, A. D., Kauppinen, T., Marinacci, C., Martikainen, P., Regidor, E., Stafford, M., and Valkonen, T. (2005). Neighbourhood unemployment and all cause mortality: a comparison of six countries. *Journal of Epidemiology and Community Health*, 59(3):231–237.

Vives, A., Amable, M., and et al., M. F. (2013). Employment precariousness and poor mental health: Evidence from spain on a new social determinant of health. *Journal of Environmental and Public Health*, 2013:1–10.

Vives, A., Amable, M., Ferrer, M., Moncada, S., Llorens, C., Muntaner, C., Benavides, F. G., and Benach, J. (2010). The employment precariousness scale (epres): psychometric properties of a new tool for epidemiological studies among waged and salaried workers. *Occupational and environmental medicine*, 67(8):548–555.

Waenerlund, A.-K., Virtanen, P., and Hammarström, A. (2011). Is temporary employment related to health status? analysis of the northern swedish cohort. *Scandinavian Journal of Public Health*, 39(5):533–539.

Wooldridge, J. M. (2010). *Econometric analysis of cross section and panel data*. The MIT press.

Table 1: Sample descriptive statistics of Self Reported Health Status (%) by employment status, year, gender, age, area of residence, family income, family wealth, and regional unemployment rate.

	Very poor	**Poor**	**Fair**	**Good**	**Excellent**	**Total**
Total	0.21	1.46	7.34	53.93	37.06	100
Employment status						
Permanent	0.13	0.91	5.77	55.24	37.95	48.58
Temporary	0.07	1.20	7.47	54.21	37.06	7.80
Unemployed	0.24	2.19	7.42	52.50	37.65	9.02
Inactive	0.34	2.11	9.50	52.41	35.65	34.60
Year						
2006	0.18	1.58	7.87	54.47	35.91	33.40
2008	0.28	1.51	7.22	53.86	37.13	33.36
2010	0.17	1.28	6.93	53.47	38.14	32.24
Gender						
Female	0.20	1.50	7.80	55.32	35.17	50.75
Male	0.22	1.41	6.87	52.50	39.00	49.25
Age						
15–34	0.06	0.30	1.76	42.31	55.58	33.50
35–54	0.18	1.36	7.08	58.65	32.73	44.66
55–64	0.50	3.44	16.44	62.11	17.51	21.84
Area of residence						
North	0.18	1.15	6.38	47.81	44.48	43.05
Center	0.25	1.37	8.20	59.42	30.76	19.48
South	0.22	1.87	8.00	58.11	31.81	37.47
Income						
$\leq 22,100$	0.34	2.82	10.45	56.56	29.83	24.97
$(22,100; 33,500]$	0.30	1.37	7.67	54.60	36.06	24.97
$(33,500; 47,900]$	0.14	1.00	6.50	53.02	39.34	25.01
$> 47,900$	0.05	0.65	4.76	51.56	42.98	25.00
Wealth						
$\leq 39,000$	0.27	2.09	8.58	54.77	34.30	25.05
$(39,000; 185,000]$	0.30	1.85	7.92	55.22	34.71	24.87
$(185,000; 330,000]$	0.20	0.97	6.64	52.78	39.41	24.82
$> 330,000$	0.06	0.94	6.24	52.97	39.79	25.26
Unemployment rate						
≤ 4.3	0.20	1.30	7.21	48.08	43.21	25.24
$(4.3; 5.7]$	0.20	1.34	6.58	53.90	37.99	20.16
$(5.7; 12.6]$	0.22	1.49	7.90	54.89	35.50	32.67
> 12.6	0.21	1.72	7.37	59.27	31.44	21.93

Table 2: Fixed-Effects (FE) and Random-Effects (RE) Ordered Logit Models for Self Reported Health Status

	M1		M2	
	FE	RE	FE	RE
	coeff. (s.e.)	coeff. (s.e.)	coeff. (s.e.)	coeff. (s.e.)
Ref: Permanent				
Temporary	**-0.149** (.102)	**-0.437** (.058)	**-0.238** (.164)	**-0.523** (.097)
Unemployed	**-0.370** (.116)	**-0.618** (.061)	**-0.506** (.163)	**-0.640** (.101)
Inactive	**-0.278** (.095)	**-0.366** (.038)	**-0.342** (.114)	**-0.410** (.056)
2008	**-0.133** (.042)	**-0.080** (.032)	**-0.226** (.062)	**-0.137** (.046)
2010	**-0.270** (.083)	-0.003 (.035)	**-0.294** (.100)	-0.018 (.049)
Fam. Income	0.035 (.032)	**0.131** (.011)	0.036 (.032)	**0.131** (.011)
Fam. Income sq.	**-0.004** (.001)	**-0.002** (.000)	**-0.004** (.001)	**-0.003** (.000)
Fam. Wealth	**0.005** (.001)	**0.001** (.000)	**0.005** (.001)	**0.001** (.000)
Fam. Wealth sq.	**-0.000** (.000)	**-0.000** (.000)	**-0.000** (.000)	**-0.000** (.000)
Unemp. rate	**0.103** (.041)	**-0.058** (.004)	**0.104** (.041)	**-0.059** (.005)
Age in 2006		**-0.085** (.002)		**-0.085** (.002)
Temp. $\times$ 2008			0.198 (.189)	0.122 (.131)
Temp. $\times$ 2010			0.052 (.205)	0.130 (.132)
Unem. $\times$ 2008			**0.368** (.167)	0.130 (.130)
Unem. $\times$ 2010			0.034 (.184)	-0.050 (.128)
Inact. $\times$ 2008			**0.135** (.098)	0.106 (.072)
Inact. $\times$ 2010			0.054 (.113)	0.029 (.002)
Hausman test	$Pr\left(68.05 > \chi^2_7\right) = 0.00$		$Pr\left(79.12 > \chi^2_{13}\right) = 0.00$	
Log-lik	-4,095.70	-32,008.29	-4,091.06	-32,005.77
Observations	37,477	37,477	37,477	37,477

Coefficients statistically significant at the 10% level are in bold.

Table 3: Diagnostic tests for time differences in Self Reported Health Status by employment status

	$H_0 : \beta_j = \gamma_{2008} + \phi_{j,2008}$		$H_0 : \beta_j = \gamma_{2010} + \phi_{j,2010}$	
	$\delta_{j,2008}$	$s.e.(\delta_{j,2008})$	$\delta_{j,2010}$	$s.e.(\delta_{j,2010})$
Whole sample				
Temporary	**-0.347**	0.170	-0.012	0.160
Unemployed	-0.216	0.176	**-0.703**	0.253
Inactive	**-0.399**	0.242	**-0.710**	0.211

Coefficients statistically significant at the 10% level are in bold.

Table 4: Fixed-Effects ordered logit models for Self Reported Health Status by gender

	Male		Female	
	M1	M2	M1	M2
	coeff. (s.e)	coeff. (s.e)	coeff. (s.e)	coeff. (s.e)
Ref: Permanent				
Temporary	**-0.352** (.149)	**-0.508** (.241)	0.088 (.163)	0.105 (.253)
Unemployed	**-0.692** (.166)	**-0.859** (.229)	-0.039 (.189)	-0.162 (.265)
Inactive	**-0.431** (.165)	**-0.440** (.193)	-0.051 (.157)	-0.124 (.180)
2008	**-0.171** (.062)	**-0.211** (.081)	**-0.112** (.062)	**-0.222** (.102)
2010	**-0.213** (.124)	**-0.266** (.139)	**-0.274** (.120)	**-0.273** (.157)
Temp. × 2008		0.313 (.271)		0.003 (.291)
Temp. × 2010		0.108 (.302)		-0.065 (.304)
Unem. × 2008		**0.390** (.230)		**0.339** (.262)
Unem. × 2010		0.125 (.253)		-0.023 (.295)
Inact. × 2008		-0.088 (.158)		**0.183** (.139)
Inact. × 2010		0.139 (.192)		0.009 (.159)
Hausman test	Rej.	Rej.	Rej.	Rej.
Log-lik	-1,859.27	-1,854.65	-1,869.03	-1,866.66
Observations	18,458	18,458	19,019	19,019

Coefficients statistically significant at the 10% level are in bold. Model specification also include family income, family income square, family wealth, family wealth square, regional unemployment rate. The Hausman test rejects the null hypothesis at the 5% level. Estimation results of the RE model are not presented here for brevity and are available upon request.

Table 5: Diagnostic tests for time differences in Self Reported Health Status by employment status: gender

	$H_0 : \beta_j = \gamma_{2008} + \phi_{j,2008}$		$H_0 : \beta_j = \gamma_{2010} + \phi_{j,2010}$	
	$\delta_{j,2008}$	$s.e.(\delta_{j,2008})$	$\delta_{j,2010}$	$s.e.(\delta_{j,2010})$
Male				
Temporary	**-0.597**	0.252	-0.297	0.237
Unemployed	**-0.594**	0.248	**-1.169**	0.354
Inactive	-0.552	0.366	**-0.830**	0.307
Female				
Temporary	0.025	0.260	0.327	0.245
Unemployed	0.104	0.282	-0.166	0.398
Inactive	-0.065	0.363	-0.463	0.335

Coefficients statistically significant at the 10% level are in bold.

Table 6: Fixed-Effects ordered logit models for Self Reported Health Status by age

	15–34		35–54		55–64	
	M1	M2	M1	M2	M1	M2
	coeff. (s.e)	coeff. (s.e)	coeff. (s.e)	coeff. (s.e)	coeff. (s.e)	coeff. (s.e)
Ref: Permanent						
Temporary	**-0.208** (.161)	**-0.475** (.269)	**-0.222** (.171)	0.147 (.265)	-0.050 (.302)	-0.202 (.487)
Unemployed	**-0.357** (.175)	**-0.603** (.249)	**-0.592** (.223)	**-0.648** (.334)	**-0.600** (.370)	**-0.548** (.583)
Inactive	**-0.275** (.180)	**-0.309** (.221)	**-0.292** (.176)	**-0.344** (.220)	**-0.534** (.188)	**-0.591** (.225)
2008	-0.096 (.079)	**-0.349** (.151)	**-0.168** (.066)	**-0.184** (.080)	-0.108 (.091)	**-0.227** (.164)
2010	-0.038 (.155)	-0.030 (.227)	**-0.482** (.132)	**-0.477** (.147)	-0.079 (.187)	-0.089 (.350)
Temp. × 2008		0.354 (.322)		0.045 (.290)		0.246 (.611)
Temp. × 2010		0.407 (.359)		-0.296 (.322)		0.252 (.618)
Unem. × 2008		**0.582** (.255)		0.331 (.330)		0.253 (.602)
Unem. × 2010		0.151 (.283)		-0.178 (.375)		-0.320 (.713)
Inact. × 2008		**0.275** (.205)		-0.057 (.176)		0.167 (.206)
Inact. × 2010		-0.228 (.244)		0.150 (.207)		0.005 (.244)
Hausman test	Rej.	Rej.	Rej.	Rej.		
Log-lik	-1,043.50	-1,036.65	-1,692.63	-1,688.37	-870.21	-869.14
Observations	12,555	12,555	16,737	16,737	8,185	8,185

See notes to Table 4

Table 7: Diagnostic tests for time differences in Self Reported Health Status by employment status: age

	$H_0 : \beta_j = \gamma_{2008} + \phi_{j,2008}$		$H_0 : \beta_j = \gamma_{2010} + \phi_{j,2010}$	
	$\delta_{j,2008}$	$s.e.(\delta_{j,2008})$	$\delta_{j,2010}$	$s.e.(\delta_{j,2010})$
15-34				
Temporary	**-0.508**	0.282	-0.124	0.250
Unemployed	**-0.580**	0.273	**-0.957**	0.440
Inactive	-0.724	0.457	**-0.892**	0.376
35-54				
Temporary	-0.316	0.276	0.037	0.260
Unemployed	-0.169	0.352	-0.694	0.430
Inactive	-0.046	0.402	**-0.675**	0.405
55-64				
Temporary	-0.205	0.494	0.025	0.483
Unemployed	-0.463	0.594	-0.793	0.866
Inactive	-0.847	0.685	-0.844	0.671

Coefficients statistically significant at the 10% level are in bold.

Table 8: Fixed-Effects ordered logit models for Self Reported Health Status by area of residence

	North		Rest of Italy	
	M1	M2	M1	M2
	coeff. (s.e.)	coeff. (s.e.)	coeff. (s.e.)	coeff. (s.e.)
Ref: Permanent				
Temporary	-0.113 (.168)	-0.121 (.259)	**-0.166** (.129)	**-0.371** (.212)
Unemployed	**-0.474** (.229)	**-1.068** (.397)	**-0.352** (.140)	**-0.518** (.190)
Inactive	-0.170 (.147)	**-0.249** (.175)	**-0.351** (.126)	**-0.438** (.153)
2008	0.069 (.067)	0.011 (.090)	**-0.305** (.056)	**-0.481** (.087)
2010	-0.238 (.200)	**-0.300** (.209)	**-0.292** (.092)	**-0.319** (.119)
Temp. × 2008		-0.164 (.305)		**0.509** (.243)
Temp. × 2010		0.198 (.321)		0.064 (.264)
Unem. × 2008		**0.709** (.437)		**0.505** (.190)
Unem. × 2010		**0.776** (.445)		0.009 (.211)
Inact. × 2008		0.144 (.150)		**0.204** (.131)
Inact. × 2010		0.134 (.177)		0.055 (.150)
Hausman test	Rej.	Rej.	Rej.	Rej.
Log-lik	-1,719.02	-1,715.23	-2,359.46	-2,351.26
Observations	16,132	16,132	21,345	21,345

See notes to Table 4

Table 9: Diagnostic tests for time differences in Self Reported Health Status by employment status: area of residence

	$H_0 : \beta_j = \gamma_{2008} + \phi_{j,2008}$		$H_0 : \beta_j = \gamma_{2010} + \phi_{j,2010}$	
	$\delta_{j,2008}$	$s.e.(\delta_{j,2008})$	$\delta_{j,2010}$	$s.e.(\delta_{j,2010})$
North				
Temporary	-0.292	0.278	-0.131	0.253
Unemployed	**-0.773**	0.440	**-0.904**	0.498
Inactive	-0.452	0.375	**-0.958**	0.476
Rest of Italy				
Temporary	**-0.384**	0.218	0.110	0.207
Unemployed	-0.202	0.202	**-1.027**	0.316
Inactive	-0.505	0.319	**-0.943**	0.259

Coefficients statistically significant at the 10% level are in bold.