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DISCUSSION PAPER SERIES

IZA DP No. 13116

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ABSTRACT

Is Precarious Employment Bad for Worker Health? The Case of Zero Hours Contracts in the UK*

The increasing numbers of workers in employment with little to no job security, so-called precarious employment, has led to a range of concerns over worker outcomes. A particular focus is the effect of instability on health in general, and particularly, mental health. We provide new evidence on this, focusing on an extreme form of precarious employment that has grown rapidly in the UK, zero-hours contracts (ZHCs). We demonstrate that workers employed on ZHCs are more likely to report a long-lasting health problem than workers employed on other types of contract. In particular, reported levels of mental ill health are higher (almost double) among ZHC workers than for other workers. These associations remain, and a positive association between ZHC employment and physical ill health emerges, after controlling for a rich set of observable characteristics. Estimated associations vary little between different demographic groups, although they are concentrated in parts of the economy where underlying job instability is likely to be higher. Finally, we exploit sectoral variation in the historical prevalence of ZHC-like employment, in an instrumental variables framework, to demonstrate large and potentially causal effects of ZHC employment on reporting a long-lasting health problem and on mental ill health, but no effect on physical health. It is unlikely that these effects are currently factored into short-term employment conditions or that they attract compensating wage differentials.

JEL Classification: J21, J48, M55

Keywords: zero hours contracts, atypical employment, precarious employment, casual employment, health, mental health, instrumental variables

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1. Introduction

Recent years have seen growing concern about the rise of precarious employment – jobs with little or no job security, often coupled with low levels of pay and working-hours fluctuations – and its effects on worker wellbeing (e.g. Kalleberg, 2009). A key question is whether such jobs impact negatively on worker health, and in particular on mental health. One potential mechanism for precarious employment to affect health is the adverse impact of job insecurity on psychological stress (Cheng and Chan, 2008; Ferrie, 2001; Green, 2015; Sverke et al., 2002), which could be exacerbated by working-hours fluctuations over which the worker has limited control. Higher-levels of sickness-related presenteeism among workers in precarious jobs, including where such jobs offer limited entitlement to paid sick leave, is another. Workers in precarious jobs may also be more at risk of work-related injuries and illness given less exposure to occupational health and safety training, less familiarity with work environments and practices, and differential assignment of tasks where there is greater exposure to work hazards (Aronnson, 1999; Benavideds et al., 2006; Green, 2015). On the other hand, the flexibility afforded by some (but as we discuss below, not all) precarious jobs could potentially drive a positive effect on health for some workers, or enable some workers with existing health problems to work where that might not otherwise be possible.

This paper examines this question specifically for the case of zero hours contracts (ZHCs) in the UK, a form of employment which has experienced a dramatic increase in prevalence over the past decade (Farina et al., 2020), and is found not only in the UK but much more widely across countries (O'Sullivan, 2019). Because ZHCs in the UK offer essentially zero job security (employers are under no contractual obligation to offer ZHC workers any hours of work at all) (Adams and Prassl, 2018), because they may feature fluctuating working hours outside of the worker's control (Low Pay Commission, 2018), and because they are typically low paid (Koumenta and Williams, 2019), they can be thought of as an extreme form of precarious employment. It follows, therefore, that if there is an adverse effect of precarious employment on workers' health, and in particular on workers' mental health, you'd expect to see it for ZHCs in the UK. ZHCs are also a live policy issue in the UK and more widely (e.g. DBEIS, 2019; O'Sullivan, 2019), but these policy debates are taking place against a background of very little existing quantitative evidence on the health (or indeed other) impacts of ZHCs on workers.

As well as addressing a particular gap in the empirical literature on ZHCs, our research fits into the wider literature examining the link between different types of employment contract and workers' health. In particular, extensive earlier work from a variety of international settings demonstrates largely negative associations between temporary or atypical employment (variously defined) and mental or general health (for reviews see Quinlan et al., 2001; Virtanen et al., 2005). Because of non-random selection of workers into employment types, however, it is not generally possible to interpret these associations as demonstrating a causal effect of employment type on health. A more recent literature attempts to address this, most commonly by exploiting longitudinal data to provide within-worker estimates that are more plausibly interpretable as approaching causal estimates. For the UK, Bardasi and Francesconi (2004) and Robone et al (2011) find no effect of temporary employment on mental or general health. For casual employment in Australia, which shares the key 'no-guaranteed hours' characteristic of ZHCs in the UK (Farina et al., 2020), Richardson et al. (2012) and LaMontagne et al. (2014) similarly find no evidence of an adverse effect of casual employment on mental health. While these approaches mitigate some concerns regarding reverse causation and sorting into employment contracts on time-invariant unobservables, they do nothing to mitigate potential concerns over remaining time-varying confounders (e.g. unobserved changes in family circumstances) or concerns regarding measurement error in contractual status, which together could bias the estimate of interest in an uncertain direction. Recognising this, Moscone et al. (2016) instruments temporary employment status with firm-level propensity to use temporary contracts (and other firm-level characteristics) and reports a negative effect of temporary employment on mental health in the Italian province of Lombardy. Given the earlier zeroes from the other longitudinal studies cited above, however, the weight of evidence for widespread negative health effects of temporary employment is far from compelling.

We examine the health effects of ZHCs using representative cross-sectional survey data for the UK, drawn from the Labour Force Survey (LFS), which provides detailed information on employment arrangements including type of contract, together with a range of long-term health indicators including for general, mental, and physical health conditions. Initially, we demonstrate that ZHC workers are more likely than other workers to report a long-lasting health problem and substantially more likely to report a mental health problem. These associations remain, and a positive association between ZHC employment and physical ill health emerges, after controlling for a rich set of observable individual, household and job characteristics. We show that these estimated associations vary little between different demographic groups, but are driven predominantly by private sector jobs (where one might imagine underlying job security is lower than that in the public sector) and concentrated in a limited number of sectors and occupations. Finally, because of the joint challenges of non-random selection of workers/jobs into contractual types and measurement error, we adopt an instrumental variable (IV) strategy where we exploit differences in the underlying propensity of jobs to be suitable for these contractual forms. Specifically, we instrument the likelihood of a current job being a ZHC with historical prevalence of ZHC-like jobs, including casual jobs, at a highly disaggregated sectoral level, and predating the recent rise in the prevalence of ZHCs. These IV estimates demonstrate large and potentially causal effects of ZHC employment on reporting a long-lasting health problem and on reporting mental ill health, but no effect on physical health.

2. ZHCs in the UK

Although there is no universally accepted single definition of a ZHC, even within the UK (Adams and Prassl, 2018), ZHCs have been defined by the UK government as employment contracts where the employer does not guarantee the individual any work and the individual is not obliged to accept any work offered (DBIS 2013). Evidence from the CIPD, however, and more recently from the Low Pay Commission, suggests that ZHC workers are often expected to accept work when offered (CIPD, 2015; Low Pay Commission, 2018), in which case the defining legal characteristic of a ZHC in practice is that the employer does not guarantee the ZHC worker any work. This now appears to be the ONS's (Office for National Statistics) preferred definition of a ZHC (ONS, 2018). These two characteristics – the lack of guaranteed hours and potentially fluctuating work hours and schedules at the employer's behest – define ZHCs as an extreme form of precarious employment. ZHC jobs are also typically low-paid (Koumenta and Williams, 2019). Note, however, that some ZHCs may not be poorly paid, may offer workers genuine flexibility in accepting hours of work, and may in practice be long-lasting with regular hours despite the no-guaranteed-hours clause. Also note that dropping the no worker obligation clause blurs the distinction between ZHCs and other forms of precarious employment. In particular, casual contracts share the no-guaranteed-work characteristic in that they can, in practice, be severed at any time with no notice period, as might some on-call contracts.

The growth of ZHCs observed in recent years in the UK has served to centre political and economic debate on the trade-offs associated with this type of employment. On the one hand, ZHCs may be particularly attractive for firms facing erratic and unpredictable demand or, in the cases where employers do allow for flexibility on the worker side, for workers who require more flexibility in hours compared to that offered by other working arrangements. On the other hand, there are the characteristics of ZHCs which may mark them out as poor quality jobs from a worker perspective, including the lack of job security, limited access to work-related benefits, training, entitlements and opportunity for career development, and unpredictability of hours and earnings. A wide range of policy interventions have been mooted in recent years to address some of these downsides of ZHCs. For example, the UK Government has recently consulted on increased regulation of, or compensation for ZHCs, in particular to address the one-sided flexibility issue (see DBEIS, 2019). There have even been calls to ban ZHCs altogether (e.g. Labour Party, 2019), a step already taken in New Zealand. Other countries, e.g. the Republic of Ireland, have recently increased the regulation of ZHCs, in effect converting them into short-hours contracts with some guaranteed hours, although loopholes remain (O'Sullivan, 2019).

Ultimately, how we respond to the growth in ZHCs should be informed by the impact of ZHCs on worker wellbeing, including on workers' health. Yet existing evidence on how ZHCs affect workers' health is sparse, and what little exists is mostly qualitative. These qualitative studies tend to show a perceived detrimental impact of ZHC employment on health, particularly mental health (e.g. Ball et al., 2017; Ndzi et al., 2017). This is echoed by anecdotal evidence in media reports (e.g. O'Connor, 2019). Quantitative evidence to date is particularly sparse and to some extent conflicting. On the one hand, the CIPD (2015) reports survey evidence that ZHC workers are less likely to feel overloaded and under excessive pressure than other workers, suggestive of but not explicitly demonstrating a positive association with mental health. On the other hand, Henderson (2019) shows higher levels of reported poor mental health among 25 year-olds on ZHCs than among 25 year-olds in other forms of employment, even after controlling for observable worker characteristics. As we have discussed for temporary employment, however, ZHC status is likely to be endogenous because of selection driven by unobservable characteristics correlated with health outcomes, simultaneity, and measurement error in survey data on ZHC status, all of which limit the extent to which we can interpret the Henderson estimate as causal.

3. Data

The data used in this paper are drawn from the UK Quarterly LFS. We pool together data over the period 2015-2018 and restrict analysis to individuals aged 16+ years in employment, excluding the self-employed. The LFS has collected data on ZHCs since 2000, on a biannual basis in quarter 2 (Q2) and quarter 4 (Q4), and we retain only these quarters in our analysis sample. Specifically, LFS respondents in these quarters are asked whether they are on a special working-hours contract and can choose up to three options among the following alternatives: flexitime, annualised hours contract, term time work, job-sharing, nine-day fortnight, four-and-a-half day week, zero hours contract, on-call working (only added as an option from 2011) or none of the above. ONS (2018) shows that the proportion of people in employment who report they are employed under a ZHC in their main job has grown rapidly over the last few years, from 0.5% in 2006 to 2.8% (or 901,000 workers) in 2017, approximately where it has remained since (the latest available estimate, for 2019 quarter 2, suggests 896,000 workers on ZHCs in their main job).

Our focus on the period 2015-2018 is motivated by a trade-off between reducing the scope for skewed measurement error to bias our estimates, and retaining sufficient sample size to support precise estimation. There are particular concerns regarding the accuracy of the LFS in measuring the prevalence of ZHCs in the UK labour market prior to the year 2015. According to ONS (2014), the data before 2013/14 are likely to underestimate the number of people on ZHCs because not all ZHC workers knew they were employed under a ZHC. Their conjecture is that this began to change rapidly as a result of increased media attention during 2013 and subsequently. Farina et al. (2020) attempt to quantify this effect, concluding that increased public awareness can account for between one quarter and two thirds of the observed rapid growth in reported ZHC prevalence over the period 2013-2014, with no clear relationship in subsequent years.[1] Nevertheless, even if we rule out such systematic under-reporting of ZHCs in the LFS from 2015 onwards, we cannot rule out the possibility that these contractual arrangements continue to be measured with error, including where survey information is collected via proxy interview.

Table 1 presents summary statistics for our sample separately by ZHC status. ZHC workers receive lower hourly wages than other workers, work fewer hours per week, are concentrated among younger workers, women, migrants, full-time students, in personal service and elementary occupations, and disproportionately in the distribution, accommodation and restaurant sector.

[Table 1 here]

LFS respondents report information on their health status in every quarter. We construct multiple indicators here. First, we construct a binary dummy equal to 1 if the respondent reports having 'any health condition or illness lasting (or expected to last) 12 months or more' and 0 otherwise. Second, for those reporting a long-term health condition, the LFS asks respondents to indicate what type or types of health problem they suffer from, drawn from a list of 17 options. Given small numbers of reported cases for some types, we aggregate these into six categories as follows: muscular/skeletal, sensorial, circulatory/breathing, digestive/kidneys/diabetes, mental, and other. Respondents are then asked what is their *main* health problem among those reported in the previous question. We use the six mutually exclusive binary indicators for main health condition constructed from this question, using the aggregations described above, to examine the relationship between ZHC status and health condition by type. For some analysis, we further aggregate the five categories other than mental into a single catchall non-mental category, which we also describe as physical health.

[1] A further motivation for this restriction is that, between 2004 and 2013 in the Q2 survey, LFS respondents reporting that they were engaged in shift work were routed away from the question on special working-hours contracts, leading to non-trivial undercounting relative to Q4 (Farina et al., 2020).

[Table 2 here]

Table 2 Panel A presents sample proportions reporting a long-term health condition. We report sample proportions for the whole sample (Column 1), for people in employment not on ZHCs (Column 2) and for workers on ZHCs (Column 3). Column 4 reports t-ratios on the ZHC dummy variable from simple linear regressions for each health indicator with no controls. Approximately 25% of all people in employment, and workers not in ZHC-employment, report having a long-lasting health problem. This figure increases to 30% for ZHC workers, and the difference is highly statistically significant. In Panel B, we report the equivalent sample proportions for main health problem. The mental health category immediately stands out, with approximately 6% of ZHC workers reporting a long-lasting mental health condition compared to approximately 3% of non-ZHC workers, again with the difference being highly statistically significant. For three of the other five categories – muscular/skeletal, sensorial, and other – prevalence among ZHC workers is significantly higher than among non-ZHC workers, albeit the differences are smaller than in the case of mental health. For circulatory/breathing and digestive/kidney/diabetes prevalence is lower among ZHC workers than among other workers.

As these health indicators are all self-reported, we also cannot rule out measurement error in our outcome variables. This will not impart bias to our estimates so long as any such measurement error is not skewed in a particular direction, although the precision of our estimates may be reduced. If there is skewed measurement error, however, e.g. because of systematic under-reporting of one or more health conditions due to stigma, then any adverse effects of ZHC employment on these health outcomes may be under-estimated.

4. Approach to Estimation

Our initial step is to provide a series of conditional associations between ZHCs and our range of health indicators, and as a result we estimate a series of (univariate) probit models as in Equation (1). Probits are preferred to linear probability models here because many of our health indicators are reported with low probability (see Long, 1997).

$$\Pr\left(Y_i = 1 \middle| ZHC_i, X_i, \tau_q\right) = \Phi\left(\beta \mathrm{ZHC_i} + \mathrm{X}_i'\delta + \tau_q\right) \qquad (1)$$

$\mathrm{Y_i}$ encompasses the set of binary health outcomes presented in Section 2, $\mathrm{ZHC_i}$, our regressor of interest, takes value 1 for LFS respondents on a ZHC and 0 otherwise, X_i includes a series of demographic characteristics (age group, gender, marital status, ethnic group, UK/British citizenship, highest educational qualification achieved, full-time student status, families with children in the age group 0-4 and 5-15, regional dummies), other (in some cases overlapping) contingent contractual forms (permanent agency workers and temporary job categories including casual, seasonal, fixed period, temporary agency and other residual temporary contracts) and job characteristics (a part-time job dummy (self-reported) and hours of work categories, tenure categories, and 1-digit occupation and industry indicators), while τ_q is a set of quarter/year dummies. We estimate (1) on the full sample and also, to examine evidence for heterogeneous effects, on samples split by age group, gender, and a series of other individual and job characteristics. With the partial exception of Henderson (2019), which is limited to estimating ZHC/mental health associations for 25 year olds, this paper is the first to present conditional associations between ZHC status and multiple health outcomes across all working ages, and the first to do so separately across numerous socio-demographic and job-characteristic groups.

Interpreting the average marginal effects on the ZHC dummy in (1) as causal, however, requires us to assume ZHC status is exogenous. As in the wider atypical work and health literature, there are several reasons why this assumption is highly questionable here. First, ZHC workers are likely to be different from other workers in unobservable ways, some of which may be correlated with health outcomes, just as they are in observable ways (see Table 1). If so, estimates from univariate models like (1) may be biased, and in an uncertain direction. Second, those with existing health problems may be more likely to take ZHC jobs than those without existing health problems, perhaps because they offer the flexibility required to be able to manage work while living with a long-term health condition, or perhaps because those with existing health conditions are less likely to be offered other types of work.[2] Estimates from (1) would likely over-estimate any adverse causal effect of ZHC status on health as a result. On the other hand, given fears about one-sided flexibility and 'zeroing down' (losing access to shifts if you do not accept all shifts offered to you), workers with existing health conditions may be *less* likely to take ZHC work, which could bias the estimated ZHC effect in (1) in the opposite direction. Finally, measurement error in the ZHC indicator, whether symmetric or skewed towards under-reporting, would bias estimated ZHC effects in (1) towards zero.

Combined, these potential biases act in an uncertain direction, making it difficult to even interpret estimates from (1) as lower or upper bounds. In the absence of a randomised controlled trial, which is difficult to envisage implementing in this context, we therefore complement the model above by implementing an IV strategy, instrumenting for the endogenous variable – ZHC status – with

[2] Although we are aware of no evidence on this specifically for ZHC employment, Dawson *et al.* (2015) uses British Household Panel Study data to show that workers in permanent positions who experience poor mental health are more likely to transition into temporary employment than workers who do not experience poor mental health.

historical variation in ZHC-like employment shares at the regional and sectoral levels. Because both the dependent variable and the endogenous regressor are dichotomous in nature, and further because both have low probabilities, this takes the form of a recursive bivariate probit model (see Wooldridge, 2002; Chiburis et al., 2012). Specifically, we estimate the following model:

$$ZHC_i^* = \alpha_{ZHC} + Z_i'\gamma + \tau_q + \varepsilon_{1i}$$

$$ZHC_\mathrm{i} = 1\ [ZHC_\mathrm{i}^* > 0]$$

$$Y_i^* = \alpha_Y + \pi ZHC_i + X_i'\varphi + \tau_q + \varepsilon_{2i}$$

$$Y_i = 1\ [Y_i^* > 0]$$

$$E[\varepsilon_1] = E[\varepsilon_2] = 0$$

$$Var[\varepsilon_1] = Var[\varepsilon_2] = 1$$

$$Cov[\varepsilon_1, \varepsilon_2] = \rho \qquad (2)$$

In (2), Y_i is the binary (health) outcome of interest for individual i, equal to 1 if the latent variable Y_i^* is greater than 0; ZHC_i is the endogenous ZHC indicator, equal 1 if the latent variable ZHC_i^* is greater than 0; X_i is defined as before; Z_i is a vector of observable characteristics including at least one variable (the instrument) that is not in X_i ; and ρ represents the correlation between the error terms in the two equations.

4.1 Validity of the Instrument

The variable in Z that is omitted from X – the instrument – is constructed using LFS data for the period 2009-2010 (or 2001-2010 in sensitivity analysis), otherwise following the sample inclusion criteria as set out in Section 3, and aggregates the total number of workers employed under a ZHC or casual contract – which, following Farina et al. (2020) we label no-guaranteed hours contracts (NGHCs) – as a share of all people in employment in a given 4-digit level industry, at the national level. Our identification strategy matches these historical 4-digit level industry cells of the share of NGHCs to individual i's current ZHC status.[3] This allows us to exploit a strong source of plausibly exogenous variation: workers in a given industry today are more likely to be on a ZHC if they are hired in a sector in which NGHCs have been used in the past.[4] At the same time, by exploiting a non-contemporaneous relation between the instrument and the instrumented variable, it is reasonable to

[3] Due to changes over time in the 4-digit sectoral classification used by the LFS, we are not able to exploit data prior to the year 2009 for this purpose. In sensitivity analysis, however, we do so at the 2-digit level.
[4] Moscone et al. (2016) use a similar argument to motivate their firm-level instruments, although these are contemporaneous rather than lagged, as in our case.

assume that individual i's current health problems will be affected by the historical proportion of NGHCs in a given industry only via the instrumented variable, i.e. her ZHC status.

Consider the validity of this IV approach in more detail. It requires that, even for like-for-like jobs and like-for-like workers, disaggregated industries vary in their propensity to use NGHCs. We observe the distribution of NGHCs at this disaggregated level at some point in the past (predating the dramatic growth of ZHC use in the UK since 2012), and then use this to instrument where we expect ZHCs to have appeared in the 2015-2018 period. This approach relies on there being factors that affect the suitability of disaggregated industries for NGHCs in the past which are sufficiently persistent to help predict prevalence of ZHCs in the present. Specifically, we have in mind persistent volatility in the demand for a sector's outputs. We know from CIPD (2015) that variability of demand is a key determinant of ZHC use at the firm level, suggesting potential for a strong first-stage association. For this instrument to be validly excludable, however, requires an assumption that persistent sector-level variability in demand does not impact on our measures of worker health other than through the nature of the employment contracts employed in that sector. Is this reasonable? Imagine that all workers are identical upon job-entry and that all jobs are permanent, full-time and secure, but that some are in sectors with volatile demand and others are not. Our assumption is equivalent to ruling out any difference in the particular health outcomes studied here, on average, between jobs in high variability and low variability-in-demand sectors. In the latter, workers are likely to be able to work at a steady pace. In the former, workers may have periods of intense work and periods of slower work. Although it seems possible that the periods of intense work could lead to temporarily higher levels of tiredness and/or stress, this would likely be offset over the longer run by periods of slower work, and the health outcomes we study here are all long-term health conditions.

Another potential threat to the validity of this approach is if there are longer-run impacts of historical NGHC prevalence at the disaggregated national-sectoral level on individual health that could potentially obscure the effects of current individual ZHC status on health we seek to identify (for a discussion of this argument in the context of Bartik-type instruments in migration impact studies see Jaeger et al., 2018). This seems unlikely in our case, however, given that the endogenous variable is at the individual (not the region/state/city/county) level and the instrument is at the national-sectoral level, and given that the LFS sample in 2009/10 does not overlap with the LFS sample in 2015-2018.

A related but perhaps more salient potential threat to this identification strategy is if industries vary in their working conditions in other persistent ways that impact on long-term health, such that any IV based on industry might not be validly excluded. To mitigate against this, we include a more aggregated set of industrial controls, along with occupational and other job characteristic controls, as described above. The assumption is that these more aggregated sectoral, occupational and other job characteristic controls wash out any persistent effects of disaggregated sector on health that do not work through ZHC status. Working in a part-time job in customer services in the hotel and restaurant sector may be more detrimental to health than working in a part-time job in customer services in the banking, finance and insurance sector for several reasons, but doing so in the event catering sector is more detrimental to health than doing so in the licensed restaurant sector only because ZHCs are more prevalent in the former than in the latter.

5. Results

5.1 Univariate Probit Estimates

In Table 3 we present the average marginal effects for the association between ZHC status and the health outcomes discussed in Section 2. Recognising the fact that other contingent contractual forms exist in the UK we include these as controls and leave as the omitted case workers who do not report a ZHC or any kind of temporary contractual arrangement.

[Table 3 here]

The estimates in Column (1) demonstrate that ZHC workers are more likely to report a long-lasting health problem (+2.9 percentage points, on a base of 25%, so approximately 12% higher) compared to permanent non-ZHC workers, even after conditioning on a rich set of observable controls. This effect is statistically significant at the 1% level. There is some indication of similar patterns for other contingent contracts, including seasonal, fixed-term and other temporary contracts. The association between casual work and reporting a long-lasting health condition takes the opposite sign, but is smaller in magnitude and only marginally statistically significant. Looking across Columns (2) – (5) and (7) we see no evidence of statistically significant conditional associations between ZHC status and reporting either a muscular/skeletal, sensorial, circulatory/breathing, digestive/kidneys/diabetes, or ‘other’ health condition. Column (6), however, shows a strong conditional association between ZHC status and reporting a *mental health* condition, with ZHC workers 1.3 percentage points (~40%) more likely to report suffering from such a condition than other workers, statistically significant at the 1% level. This is a large effect. It is consistent with the evidence of a similar conditional association for 25 year olds in England presented by Henderson (2019), but here we condition on job characteristics as well as worker characteristics, and our sample covers all working ages. There is little evidence that other forms of contingent work are associated with variations in mental health, with only fixed term contracts having a small positive association with reporting mental health problems.

Because the estimated associations between ZHC status and the five other non-mental health conditions are all positive, albeit small in magnitude and statistically insignificant on their own, we also aggregate them together into a catchall measure for reporting a main health condition that is physical in nature before we draw any firm conclusions regarding the overall conditional association between ZHC status and physical health. The resulting estimate (in Column 8) shows a now statistically significant association, taking the same sign as that between ZHC status and mental health, and of a similar magnitude (1.3 percentage points, although in this case on a base of 22%, corresponding to ~6% higher prevalence).

In Table 4 we examine the conditional associations between ZHC status and reporting a long-lasting health problem, between ZHC status and reporting that your main health problem is mental, and between ZHC status and reporting that your main health problem is in any of the other five (physical) categories, separately for different sub-groups of the population. Specifically, we split the sample by age, gender, education, aggregate industry, occupation, citizenship/migrant status and public/private sector. The patterns for reporting a mental health condition and for a long-term health condition in general are very similar. In both cases all groups show either a positive association or a non-significant one; there is no group for whom the association between ZHC status and the poor health outcome is negative. The same holds for physical health conditions, albeit estimated associations are typically smaller and there are fewer that are statistically significant at conventional levels. In what follows we concentrate primarily on the first two outcomes: reporting a long-term health condition and reporting that your main health condition relates to mental health.

[Table 4 here]

Specifically, the associations between ZHC status and poor long-term health, and between ZHC status and poor mental health are statistically significant and similar in magnitude for 18-24s, 25-34s, 35-49s but are smaller and non-significant for 50+ year olds. This may in part reflect use of ZHC jobs as part of 'winding-down' retirement strategies, and is consistent with LaMontagne et al.'s (2014) finding that casual work in Australia impacts mental health differently for older workers than for younger workers (although in that case the estimates suggest a positive effect on mental health for older workers, with no effect for younger workers). There are no differences by gender for either health outcome; both men and women show a statistically significant association between ZHC employment and poor health, and of similar magnitude. Contrast this with tentative evidence of gender differences in the impact of fixed-term employment in Australia on mental health from Richardson et al. (2012), and in the impact of temporary employment on mental health in Britain from Robone et al. (2011). We find inverted U-shaped relationships between ZHC and both health outcomes in terms of education level; those at either extreme show no conditional associations while those in the middle education categories show statistically significant associations with broadly similar magnitudes. (Again there is some existing evidence for mixed effects of temporary work on health by education level in Richardson et al. (2012) and Robone et al. (2011).) Finally, in terms of individual worker characteristics, the associations between ZHCs and the two health outcomes are larger in magnitude (and only statistically significant) for UK citizens and natives compared to non-UK citizens and migrants.

Turning to job characteristics, there is a clear public/private sector split, where we only see evidence of statistically significant ZHC health effects in the private sector not the public sector (likely to reflect greater underlying job instability in the private sector than in the public sector). Further, although small sample sizes are an issue in some cases here, negative health effects also appear to be concentrated in particular industries and occupations. In particular, negative long-term health effects of ZHCs appear to be disproportionately driven by those in the restaurant/hotel sector and those in the public administration, education and health sector. (There is also a statistically significant association with ZHC status in the banking sector, but not for mental health.) The occupations that show a statistically significant association between ZHCs and health are skilled trades (overall long-term health only), personal services (for both), and (for mental health only) administrative, process/plant/machine operatives, and elementary occupations.

5.2 IV Estimates

As discussed in Section 4, the likely endogeneity of ZHC status in Equation (1) means we cannot interpret the conditional associations between ZHC status and health outcomes presented in Tables 3 and 4 as capturing causal effects of ZHC employment on health. To generate estimates that are more plausibly interpretable as approaching causal estimates, we therefore estimate the bivariate probit (IV) model set out in Model (2), instrumenting for ZHC status with historical industry prevalence of precarious contracts, with results presented in Table 5. We present bivariate probit estimates for reporting a long-term health condition, for reporting a mental health condition, and for reporting a long-term physical health condition. As before, the estimates presented are average marginal effects.[5] Note that the first stage association between the instrument and the ZHC dummy is positive and highly statistically significant in each case (they are the same, bar a slight difference in sample between Column (1) and Columns (2) and (3)), suggesting persistent sectoral-level characteristics associated with the prevalence of NGHCs. Specifically, a 1 percentage point increase in historical

[5] We follow Greene (2012), page 716, to determine the marginal effects by computing Prob[Y =1 | ZHC =1, X] - Prob[Y =1 | ZHC =0, X].

NGHC prevalence is associated with a .51 percentage point increase in ZHC prevalence between 2015 and 2018.

[Table 5 here]

First consider Column (1), which presents IV estimates of the impact on reporting a long-term health condition of being employed under a ZHC as opposed to any other contractual form. The relevant average marginal effect, which is just outside statistical significance at the 10% level, is .040, which is quite large, but not implausibly so, compared to the overall non-ZHC proportion reporting a long-term health condition of .255. This provides the first, tentative, quantitative evidence for a potentially causal impact of ZHC employment on health to be presented in the literature, and one that is economically significant in terms of magnitude, although imprecisely estimated. The estimated magnitude of this effect is larger (by a factor of one third) than the corresponding univariate model estimate (0.029), consistent with the dominant biases in the univariate model driving the estimate towards zero. This could reflect measurement error in the ZHC indicator, simultaneity whereby workers with existing long-term health conditions are less likely to take on ZHC jobs perhaps because of fears about zeroing down if they don't accept shifts, or selection into ZHC status on unobservables that are negatively correlated with reporting a long-term health condition.[6] On the other hand, the IV estimate provides only a local average treatment effect (LATE), and although the IV is broad-brush (rather than relying on a narrowly-drawn group of compliers), and although Table 4 suggests associations between ZHC status and reporting a long-term health condition are reasonably homogenous, its magnitude is not necessarily informative about the direction of these different biases and, more importantly, not necessarily meaningful from a policy perspective.

In Column (2) we present the equivalent estimates for reporting a mental health condition. The estimated average marginal effect is almost identical to that in Column (1), at .041, but in this case is statistically significant at the 5% level. The magnitude of this effect is very large relative to the proportion of non-ZHC workers reporting a mental health condition (.032), but again not implausibly so. As in the overall long-term health condition case, this is the first time a plausibly causal impact of ZHC employment on mental health has been presented in the literature, and it is also one of the first estimates in the wider contingent employment and health literature, alongside Moscone et al. (2016), to show a plausibly causal negative effect of contingent employment on mental health. As in the long-term health condition case, the fact that this estimate is larger than the corresponding

[6] The estimated correlation ρ between the errors in (2) takes a negative sign, although it is small in magnitude and statistically insignificant.

univariate estimate (.013) suggests the dominant biases in the univariate model might act in a downwards direction, although again bear in mind this is a LATE.[7]

In contrast, the final column of Table 5 suggests no adverse effect of ZHC employment on physical health; the estimate is small in magnitude, nowhere near statistically significant, and takes a negative sign.[8] Further, because the proportion of those reporting either a mental or physical health condition sums to the proportion reporting a long-term health condition overall, the suggestion is that the tentative effect of ZHC employment on reporting a long-term health condition, just outside conventional levels of statistical significance, might be driven entirely by the effect of ZHC employment on mental health. Given that there are plausible mechanisms for ZHC employment to impact on physical health, as discussed earlier in the paper, this result is interesting in its own right. But it also lends additional support to our IV strategy; to reject the validity of the IV in this case is to conjecture validity issues that are unique to *mental* health problems.

5.3. Sensitivity Analysis

We subject these conclusions to a number of robustness tests and extensions, which, taken together, provide no evidence that leads us to question our existing conclusions. First, we include the self-employed in the analysis sample, with a self-employed dummy variable included in the set of atypical employment types, and our conclusions remain unchanged.[9] Second, we further restrict the sample to 2016-2018 only, and again our key conclusions remain unchanged although estimates are less precise.[10] Third, we repeat the bivariate probit estimation using several different variants of our IV approach, and again our conclusions remain unchanged, although in some cases the estimated ZHC effect on reporting a long-term health condition becomes statistically significant at conventional levels. Results are presented in Table 6.

[Table 6 here]

In the first of these IV sensitivity analyses we retain the same NGHC instrument but control for sectoral variation in other work-related characteristics at the 2-digit rather than the 1-digit level. In the second, we construct our instrument at the 2-digit rather than the 4-digit level (controlling for other sectoral variation at the 1-digit level), which enables us to use data over the full 2001-2010

[7] In this case the estimated correlation ρ between the errors in (2) also takes a negative sign, but is larger in magnitude and statistically insignificant at the 95% level.

[8] In this case the estimated correlation ρ between the errors in (2) also takes a positive sign, but is again small and statistically insignificant.

[9] When we re-estimate Equation (1) in this case, self-employment is conditionally positively associated with reporting a long-term health condition and with reporting a mental health condition, but with magnitudes smaller in each case than the ZHC associations.

[10] In both these cases results are available from the authors on request.

period for its construction. In the third, we use the change in NGHC prevalence at the 2-digit industry level over the period 2001-2010 rather than the levels. In the fourth, we instrument for ZHCs using 4-digit NGHC shares for 2009/10 at the regional rather than the national level. In the fifth, we repeat the baseline IV model but using only the casual share, rather than the NGHC share, in each sector to construct the instrument. In the sixth, we repeat the 2-digit 2001-2010 version of the IV, again restricted to the casual share rather than the NGHC share in each sector.

6. Conclusions

Increases in precarious employment have led to growing concern about the implications of such employment for worker wellbeing, including workers' health. This paper focuses on the health effects of one particular and extreme form of precarious contract in the UK, so-called ZHCs, which have experienced a dramatic increase in prevalence over the past decade. We show that ZHC employment is associated with reporting a long-term health condition, reporting a physical health condition, and reporting a mental health condition, even after conditioning on a wide range of controls one might think influence both contractual status and health outcomes. We then go beyond this to provide the first estimates of ZHC health effects that can be credibly interpreted as causal. These IV estimates demonstrate substantial effects of ZHC employment on long-term health, driven by adverse impacts on mental health, but no effect on physical health. The implication is that increased precarious employment in the form of ZHCs has led to increased ill health among workers, likely generating health care burdens for society and, given these effects are not yet well known and the impacts are on longer-term health outcomes, seem unlikely to generate compensating differentials for individuals. Moves to regulate ZHCs in the UK and elsewhere should be seen in this light.

Furthermore, these results make a clear contribution to the wider international evidence base on the adverse health effects of atypical employment; these effects do exist and seem to be of non-trivial magnitude, at least for this particular contract form in this particular context. Our conjecture for why we find adverse health effects of ZHC employment where most others before us have found none for other forms of temporary employment is first, that ZHC employment is particularly precarious, and second, that our empirical approach addresses biases that individual fixed or quasi-fixed approaches do not.

Although there are numerous aspects of ZHC employment which may generate adverse causal effects on mental health, and although qualitative evidence exists supporting the existence of at least some of these mechanisms in the context of ZHCs, data limitations mean we are not able here to assess which of these mechanisms is most important in driving the overall effect. This is an important limitation in terms of offering specific policy implications. We cannot say, for example, whether regulating ZHCs with respect to one-sided flexibility (as suggested in DBEIS, 2019) would reduce their detrimental impacts on mental health, although this seems likely. Nor can we say whether compensating ZHC workers for uncertain hours with a higher minimum wage (as suggested by Taylor et al., 2017) would help to reduce their detrimental health effects. Banning ZHCs may displace workers onto other contingent contract forms which may also have detrimental health effects.

Nevertheless, if employers see that ZHCs can impact negatively on their workers' mental health, one imagines that many will want to act as far as possible to either improve the terms and/or implementation of those ZHCs, to provide access to counselling and support services to workers who may be affected, or even to move away altogether from employing workers under these contracts. In all three cases government can potentially play a role in supporting employers who take such steps and limiting the extent to which they can be undercut by employers who do not. The role of trade unions here is also likely to be important, including through informing workers of the health risks involved in ZHC work, which in turn could help to drive the emergence of compensating differentials. The evidence presented here also suggests particular industries and occupations where such interventions might initially be targeted.

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Table 1: Descriptive Statistics by ZHC Status

	Employed not on a ZHC	Employed on a ZHC
	Mean	**Mean**
	(St. Dev.)	**(St. Dev.)**
HOURPAY (2017£)	14.59	9.19
	(9.61)	(7.42)
HRRATE (2017£)	9.99	8.69
	(8.88)	(4.81)
Working Hours	34.017	23.689
	(11.04)	(13.61)
Permanent Agency Contract	0.014	0.047
Temporary: Agency	0.009	0.091
Temporary: Causal	0.009	0.150
Temporary: Seasonal	0.003	0.028
Temporary: Fixed Period	0.025	0.046
Temporary: Other	0.007	0.080
Age Group (16-24)	0.122	0.376
Age Group (25-34)	0.233	0.186
Age Group (35-49)	0.341	0.187
Age Group (50-64)	0.275	0.202
Age Group (65+)	0.029	0.048
Female	0.491	0.559
Marital Status: Divorced	0.072	0.062
Marital Status: Married	0.504	0.304
Marital Status: Other	0.015	0.017
Marital Status: Separated	0.025	0.029
Marital Status: Single	0.384	0.588
Children (0-4)	0.152	0.111
Children (5-15)	0.275	0.255
Non-UK/British Citizenship	0.136	0.164
Ethnic Group: Asian	0.055	0.051
Ethnic Group: Black	0.028	0.058
Ethnic Group: Chinese	0.005	0.003
Ethnic Group: Other	0.025	0.038
Ethnic Group: White	0.888	0.850
Full-time Student	0.034	0.209
Education: Degree or equiv.	0.351	0.206
Education: Higher Education	0.095	0.091
Education: GCE A level	0.223	0.302
Education: GCSE A-C	0.193	0.240
Education: Other	0.076	0.098
Education: No Qualification	0.061	0.063
Part-Time	0.254	0.671
Temporary Job	0.051	0.363
Public Employment	0.265	0.155
Tenure: (0-11) months	0.169	0.391
Tenure: (12-23) months	0.116	0.196
Tenure: (24-35) months	0.087	0.109
Tenure: (36-47) months	0.069	0.078
Tenure: (48-59) months	0.054	0.050
Tenure: 60+ months	0.505	0.176
Occup: Managers & Senior Off.	0.100	0.018
Occup: Professional	0.219	0.071

Occup.: Associate Prof. & Tech.	0.141	0.060
Occup: Admin. & Secretarial	0.119	0.058
Occup: Skilled Trades	0.079	0.047
Occup: Personal Service	0.091	0.231
Occup: Sales & Customer Serv.	0.085	0.088
Occup: Process, Plant, Mach. Op.	0.060	0.077
Occup: Elementary	0.106	0.350
Industry: Agri & Fish	0.007	0.004
Industry: Bank, Fin. & Insur.	0.164	0.103
Industry: Construction	0.051	0.016
Industry: Distrib., Hotels & Rest.	0.190	0.346
Industry: Energy & Water	0.019	0.005
Industry: Manufacturing	0.102	0.047
Industry: Other Services	0.046	0.102
Industry: Publ. Ad., Educ, Health	0.332	0.317
Industry: Transport & Comm.	0.089	0.060
Region: East Midlands	0.072	0.081
Region: Eastern	0.097	0.085
Region: London	0.135	0.114
Region: North East	0.038	0.046
Region: North West	0.108	0.105
Region: Northern Ireland	0.030	0.011
Region: Scotland	0.086	0.080
Region: South East	0.140	0.147
Region: South West	0.084	0.106
Region: Wales	0.045	0.051
Region: West Midlands	0.085	0.091
Region: Yorkshire & Humberside	0.079	0.083
N	235,139	6,625

Notes: Each entry reports the (weighted) means/proportions and standard deviation (in parentheses) for each demographic and job characteristic for all people in non-ZHC employment (column 1) and for all those in ZHC employment (column 2), in each case excluding the self-employed, pooling together Q2 and Q4 from 2015-2018 LFS. Two measures of hourly wages are presented, i.e. HOURPAY and HRRATE. The former is a derived variable constructed by ONS dividing the weekly earnings by the sum of the weekly usual working hours (excluding overtime) and the weekly usual hours of overtime. The latter reports the hourly wage for LFS respondents reporting to be paid on an hourly basis.

Table 2: Reported Health Problems by ZHC Status

	Total Sample	Employed not on a ZHC	Employed on a ZHC	t-ratio for difference
Panel A: Health Status				
Long Lasting Health Problem	0.251	0.255	0.295	191.23***
N	296,239	235,139	6,625	
Panel B: Main Health Problem				
Muscular/Skeletal	0.060	0.061	0.067	48.15***
Sensorial	0.014	0.014	0.017	47.99***
Circulatory/Breathing	0.064	0.065	0.062	-30.49***
Digestive/Kidney/Diabetes	0.036	0.037	0.035	-20.27***
Mental	0.033	0.032	0.058	240.54***
Other	0.044	0.045	0.056	102.76***
N	294,436	233,685	6,573	

Notes: Each entry reports the (weighted) proportions reporting having a long-lasting health problem (panel A) and main health problem (Panel B), obtained using the QLFS Q2 and Q4 samples in employment pooled over the period 2015-2018. Column (1) refers to all individuals in employment, excluding the self-employed. Column (2) refers to all individuals in employment, excluding self-employed, not on a ZHC. Column (3) refers to individuals in employment, excluding self-employed, on a ZHC. Sample sizes for columns (2) and (3) do not sum to the total sample for column (1) because of missing values for the ZHC indicator. Column (4) reports t-ratios for the ZHC dummy in simple linear regressions, with no additional controls, for each health indicator.

Table 3: Probit (AME) - ZHCs and Health Status

		Main Health Problem						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Long Last Health Problem	**Muscular**	**Sensory**	**Circul./Breath.**	**Digestive/ Kidney/ Diabetes**	**Mental/ Depression**	**Other**	**All Non-Mental**
ZHC	**0.029*****	0.005	0.001	-0.000	0.002	**0.013*****	0.005	**0.013****
	(0.006)	(0.004)	(0.002)	(0.003)	(0.003)	**(0.003)**	(0.003)	**(0.006)**
Perm. Agency Work	-0.012	0.001	-0.002	-0.005	-0.001	-0.003	-0.001	-0.008
	(0.008)	(0.004)	(0.002)	(0.004)	(0.003)	(0.003)	(0.004)	(0.007)
Temp.: Agency Work	-0.003	0.005	-0.001	**-0.010****	0.001	0.002	0.004	-0.002
	(0.009)	(0.005)	(0.002)	**(0.005)**	(0.004)	(0.004)	(0.005)	(0.008)
Temp.: Casual	**-0.015***	-0.006	0.002	0.006	-0.006	-0.003	-0.006	-0.009
	(0.009)	(0.005)	(0.002)	(0.005)	(0.004)	(0.003)	(0.004)	(0.008)
Temp.: Seasonal	**0.030***	0.006	0.006	0.007	0.011	-0.001	0.004	**0.033****
	(0.016)	(0.010)	(0.005)	(0.009)	(0.008)	(0.005)	(0.008)	**(0.015)**
Temp: Fixed Period	**0.023*****	0.006	**0.004****	-0.000	0.004	**0.006****	0.001	**0.016*****
	(0.006)	(0.004)	**(0.002)**	(0.004)	(0.003)	**(0.002)**	(0.003)	**(0.006)**
Temp.: Other	**0.021****	**0.016****	0.004	-0.003	0.003	0.004	0.001	**0.019***
	(0.010)	**(0.006)**	(0.003)	(0.006)	(0.005)	(0.004)	(0.005)	**(0.010)**
Demo. Characteristics	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Job Characteristics	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Regional Dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Quarter Dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	232,868	232,387	232,387	232,387	232,387	232,387	232,387	232,387
Pseudo R^2	0.045	0.048	0.010	0.037	0.033	0.052	0.024	0.051

Notes: Significance at the 10% level is represented by *, at the 5% level by **, and at the 1% level by ***. Entries are average marginal effects with associated robust standard errors in parentheses. The dependent variables are binary indicators taking value 1 if LFS respondents report long-lasting health problems (Column 1), or that their main health problem is in one of the six aggregated categories (Columns 2-7), or that their main health problem is non-mental, i.e. any one of muscular, sensory, circulatory/breathing, digestive or other (Column 8). Demographic characteristics controls are age, gender, marital status, binary indicators for the presence of children in the household, non-UK/British Citizenship, ethnic group, full-time student status, and highest qualifications achieved. Job characteristics are part-time job, public employment, tenure, occupation and (1-digit) industry indicators, and categorised number of hours worked. Regional and year/quarter dummies are also included. The estimates were obtained using the QLFS Q2 and Q4 surveys pooled over 2015-2018, retaining all people in employment, excluding the self-employed.

Table 4: Probit (AME) - ZHCs and Health, Heterogeneous Effects

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	LONG TERM HEALTH CONDITION			MAIN CONDITION MENTAL			MAIN CONDITION PHYSICAL		
	AME	s.e.	N	AME	s.e.	N	AME	s.e.	N
PANEL 1: AGE									
16-24	**0.025*** **	**0.010**	**24,260**	**0.020*** **	**0.005**	**24,230**	0.003	0.008	24,241
25-34	**0.038*** **	**0.013**	**48,099**	**0.019*** **	**0.007**	**48,044**	0.015	0.012	48,044
35-49	**0.036*** **	**0.013**	**83,284**	**0.015** **	**0.006**	**83,136**	0.018	0.012	83,136
50+	0.018	0.012	77,218	0.005	0.004	76,959	0.012	0.012	76,959
PANEL 2: GENDER									
Male	**0.024*** **	**0.009**	**112,687**	**0.012*** **	**0.003**	**112,497**	0.008	0.009	112,497
Female	**0.027*** **	**0.008**	**120,181**	**0.015*** **	**0.004**	**119,890**	0.011	0.008	119,890
PANEL 3: EDUCATION									
Degree	0.013	0.013	78,538	-0.000	0.005	78,381	0.013	0.012	78,381
Higher Education	**0.068*** **	**0.021**	**23,137**	**0.021** **	**0.010**	**23,085**	**0.046** **	**0.020**	**23,085**
GCE A Level	0.016	0.012	52,136	**0.013** **	**0.005**	**52,025**	-0.002	0.011	52,025
GCSE Level	**0.046*** **	**0.013**	**46,800**	**0.025*** **	**0.006**	**46,701**	0.017	0.012	46,701
Other Education	**0.040** **	**0.020**	**18,019**	**0.015* **	**0.008**	**17,826**	0.018	0.019	17,971
No Education	0.008	0.024	14,238	0.005	0.008	14,040	-0.002	0.023	14,224
PANEL 4: INDUSTRY									
Agri/Fish	0.007	0.086	1,526	-	-	**765**	0.037	0.089	1,523
Banking	**0.049*** **	**0.019**	**36,175**	0.005	0.007	36,105	**0.043** **	**0.018**	**36,105**
Construction	-0.021	0.039	11,442	0.010	0.014	10,560	-0.032	0.037	11,428
Restaurants/Hotel	**0.035*** **	**0.011**	**44,031**	**0.020*** **	**0.005**	**43,953**	0.010	0.010	43,953
Energy	0.115	0.096	4,446	0.022	0.039	4,033	0.104	0.093	4,440
Manufacturing	0.018	0.027	24,141	0.011	0.011	24,084	0.006	0.026	24,084
Other Services	0.023	0.021	10,631	0.002	0.008	10,571	0.021	0.020	10,611
Publ. Adm., Educ., Health	**0.027** **	**0.011**	**80,697**	**0.014*** **	**0.005**	**80,512**	0.010	0.010	80,512
Transport	-0.000	0.025	19,779	0.015	0.010	19,611	-0.019	0.023	19,731
PANEL 5: OCCUPATION									
Managers & Senior Off.	0.070	0.045	23,056	0.026	0.018	22,564	0.040	0.043	23,009
Professional	0.010	0.020	50,397	0.001	0.008	50,291	0.010	0.019	50,291
Associate Professions & Tech.	0.027	0.025	31,614	-0.002	0.009	31,418	0.027	0.024	31,553
Admin. & Secretarial	0.016	0.025	28,185	**0.029** **	**0.014**	**28,115**	-0.014	0.022	28,115
Skilled Trades	**0.047* **	**0.027**	**17,730**	-0.006	0.006	17,577	**0.054** **	**0.027**	**17,701**
Personal Service	**0.048*** **	**0.013**	**22,649**	**0.021*** **	**0.007**	**22,578**	**0.025** **	**0.013**	**22,599**
Sales & Costumer Service	0.028	0.022	19,532	0.002	0.010	19,457	0.023	0.021	19,489
Process, Plant and Machine Op.	0.012	0.021	14,223	**0.026*** **	**0.010**	**13,547**	-0.015	0.020	14,202
Elementary	0.018	0.011	25,482	**0.013** **	**0.005**	**25,428**	0.001	0.011	25,428
PANEL 6: CITIZENSHIP									
UK/British	**0.032*** **	**0.007**	**202,844**	**0.014*** **	**0.003**	**202,422**	**0.013** **	**0.006**	**202,422**
Non-UK/British	0.015	0.013	30,024	0.007	0.006	29,965	0.008	0.013	29,965
PANEL 7: MIGRANT STATUS									
UK/British	**0.033*** **	**0.007**	**197,459**	**0.015*** **	**0.003**	**197,051**	**0.014** **	**0.007**	**197,051**
Non-UK/British	0.013	0.012	35,409	0.006	0.005	35,336	0.006	0.012	35,336
PANEL 8: PUB. SECT									
Private Sector	**0.030*** **	**0.007**	**168,888**	**0.014*** **	**0.003**	**168,558**	**0.011* **	**0.006**	**168,558**
Public Sector	0.012	0.015	63,980	0.002	0.006	63,829	0.008	0.015	63,829

Notes: Significance at the 10% level is represented by *, at the 5% level by **, and at the 1% level by ***. Entries are average marginal effects (Columns 1, 4 and 7) with associated robust standard errors (Columns 2, 5 and 8) for the samples restricted to the relevant demographic group or industry/occupational category, with sample sizes given in Columns 3, 6 and 9. The dependent variables are binary indicators taking value 1 if LFS respondents report a long-lasting health problem (Columns 1-3), if their main health problem is mental (Columns 4-6) or the remaining main physical health conditions (Columns 7-9). Demographic characteristics controls are age, gender, marital status, binary indicators for the presence of children in the household, non-UK/British Citizenship, ethnic group, full-time student status, and highest qualifications achieved (excluding those where we split the sample on that dimension). Job characteristics are part-time job, public employment, tenure, occupation and (1-digit) industry indicators, and categorised number of hours worked (again excluding those where we split the sample by that dimension). Regional and year/quarter dummies are also included. The estimates were obtained using the QLFS Q2 and Q4 surveys pooled over 2015-2018, retaining all people in employment in that particular group, excluding the self-employed.

Table 5: Bivariate Probit (AME) – ZHCs and Health

	(1) LONG-LASTING HEALTH PROBLEM	(2) MAIN PROBLEM MENTAL HEALTH	(3) MAIN PROBLEM OTHER THAN MENTAL HEALTH
First Stage			
NGHCs	**0.514******	**0.514******	**0.514******
	(0.016)	**(0.016)**	**(0.016)**
N	232,849	232,368	232,368
Second Stage			
ZHC	0.040	**0.041****	-0.010
	(0.025)	**(0.016)**	(0.022)
Demographic Characteristics	Yes	Yes	Yes
Precarious Contracts	Yes	Yes	Yes
Job Characteristics	Yes	Yes	Yes
Regional Dummies	Yes	Yes	Yes
Quarter Dummies	Yes	Yes	Yes
N	232,849	232,368	232,368
ρ	-0.016	**-0.122**	0.038
Wald Test ρ=0 (p-value)	0.663	**0.040**	0.315

Notes: Significance at the 10% level is represented by *, at the 5% level by **, and at the 1% level by ***. Entries are average marginal effects with associated robust standard errors in parentheses. In the first stage we regress the binary indicator for ZHC on the instrumental variable, i.e. the share of NGHCs in a given 4-digit industry in the years 2009-2010, and controls as below. The dependent variables in the second stage are binary indicators taking value 1 if LFS respondents report a long-lasting health problem (Column 1), that their main health problem is mental (Column 2), or that their main problem is in any of the other five categories other than mental (Column 3). Demographic characteristics controls are age, gender, marital status, binary indicators for the presence of children in the household, non-UK/British Citizenship, ethnic group, full-time student status, and highest qualifications achieved. Job characteristics are part-time job, public employment, tenure, occupation and (1-digit) industry indicators, and categorised number of hours worked. Regional and year/quarter dummies are also included, as are dummies for the other contingent contract forms in Table 3. The estimates were obtained using the QLFS Q2 and Q4 surveys pooled over 2015-2018, retaining all people in employment, excluding the self-employed.

Table 6: Bivariate Probit (AME) – ZHCs and Health, Alternative Instruments

	(1)	(2)	(3)
	LONG-LASTING HEALTH PROBLEM	**MAIN PROBLEM MENTAL HEALTH**	**MAIN PROBLEM OTHER THAN MENTAL HEALTH**
Baseline (NGHCs, 4-digit, 2009/10)	0.040	**0.041****	-0.010
	(0.025)	**(0.016)**	(0.022)
NGHCs, 4-digit, 2009/10, 2-digit industry controls[a]	0.040	**0.034****	-0.006
	(0.026)	**(0.016)**	(0.024)
NGHCs, 2-digit, 2001/10	0.038	**0.043****	-0.013
	(0.026)	**(0.017)**	(0.022)
NGHCs, 2-digit change, 2001/10	**0.072****	**0.061*****	0.007
	(0.028)	**(0.020)**	(0.026)
NGHCs, 4-digit industry-region, 2009/10	**0.057****	**0.058*****	-0.007
	(0.028)	**(0.019)**	(0.025)
Casuals, 4-digit, 2009/10	0.037	**0.043****	-0.014
	(0.026)	**(0.017)**	(0.023)
Casuals, 2-digit, 2001/10	0.034	**0.043****	-0.017
	(0.026)	**(0.018)**	(0.023)

Notes: Significance at the 10% level is represented by *, at the 5% level by **, and at the 1% level by ***. Entries are average marginal effects of ZHC status on health with associated robust standard errors in parentheses. In the first stage we regress the binary indicator for ZHC on the relevant instrumental variable and controls as given below. The dependent variables in the second stage are binary indicators taking value 1 if LFS respondents report a long-lasting health problem (Column 1), that their main health problem is mental (Column 2), or that their main problem is in any of the other five categories other than mental (Column 3). Demographic characteristics controls are age, gender, marital status, binary indicators for the presence of children in the household, non-UK/British Citizenship, ethnic group, full-time student status, and highest qualifications achieved. Job characteristics are part-time job, public employment, tenure, occupation and industry indicators (at 1-digit level with the exception of [a] where industry controls are at the 2-digit level albeit with some (<10%) sectors aggregated where sample size at the 2-digit level is insufficient to support estimation), and categorised number of hours worked. Regional and year/quarter dummies are also included, as are dummies for the other contingent contract forms in Table 3. The estimates were obtained using the QLFS Q2 and Q4 surveys pooled over 2015-2018, retaining all people in employment, excluding the self-employed.