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ORIGINAL ARTICLE

Zero hours contracts and self-reported (mental) health in the UK

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Abstract

This article examines associations between precarious contract types and a range of self-reported health measures for the UK. We focus on zero hours contracts (ZHCs), an extreme form of precarious employment that has grown rapidly in the UK over the last decade, and on mental health. We demonstrate that workers employed on ZHCs are more likely to report a long-term health condition than workers employed on other types of contract, with the main driver being that they are almost twice as likely to report mental ill health. These associations survive conditioning on an extensive set of observable individual, job and contextual characteristics, and are robust to sensitivity analysis designed to explore the likely extent of bias due to unobserved confounders. We discuss potential explanations for these associations, from sorting of workers with poor health into ZHC employment to detrimental effects of ZHC employment on health, drawing on additional instrumental variables estimates to do so. Finally, we discuss potential policy implications.

1 | INTRODUCTION

Zero hours contracts (ZHCs) are a particularly precarious form of employment contract where the employer does not guarantee the individual any work and where the individual is not obliged to accept any work offered. They are increasingly prevalent in the UK and also found more widely across countries (O'Sullivan, 2019). A range of reasons have likely contributed to their increase. For instance, the decline in unionism in the UK over the past decades (Pencavel, 2004) allowed for a greater ability for employers to use flexible employment contracts in general, and this set the scene for increased use of ZHCs. At the same time, Datta et al. (2019) suggest that increased minimum wages in the UK led to employers in low-wage sectors exploring other cost margins, including the use of ZHCs. While, Farina et al. (2020) discuss the role of greater employer awareness of ZHCs as a possible contractual form as a source of their growth.

The particular form of ZHCs in the UK offers essentially zero job security because employers are under no contractual obligation to offer ZHC workers any hours of work at all (Adams & Prassl, 2018). They also often feature fluctuating working hours outside of the worker's control in practice (Low Pay Commission, 2018) and are typically low-paid (Koumenta & Williams, 2019). Given their highly precarious nature, we might expect a strong association between ZHC employment and worker health, particularly mental health, whether because ZHC employment is detrimental to health or because workers with poor health sort themselves or are sorted into ZHC employment.

There is an extensive literature showing negative associations between worker mental and occupational health and other forms of precarious employment (for reviews, see Quinlan et al., 2001, and Virtanen et al., 2005). Recently, and as surveyed in Wilson and McDaid (2022), a body of research on ZHCs has developed, that among others, is concerned with the effect ZHCs on worker health. As highlighted in this survey, many of these are from small samples or specific settings that are difficult to generalize, focus specifically on gig-workers or other specific aspects such as differential effects of contractual arrangements on mental health during Covid (Apouey et al., 2020). In general, these articles also find negative associations between ZHCs and a variety of health outcomes, including mental health. Most closely related to our article is Henderson (2019), which focuses on a cohort of 25 years olds in the UK and finds that ZHC workers are more likely to report mental health problems.

This article returns to this issue and aims to contribute to the literature by examining the associations between ZHC employment and self-reported health outcomes, including mental ill health, across the whole UK and the full working-age range, in the years prior to the covid-19 pandemic. We also compare ZHCs with other precarious contract (OPC) forms in this respect. We use the primary source of micro survey data on ZHCs for the UK: the Quarterly Labour Force Survey (QLFS). The QLFS provides detailed information, for a large nationally and demographically representative sample of working-age adults, on employment arrangements, including ZHCs, and on a range of long-term health indicators including for general, mental and physical health conditions. We first present estimated unconditional associations between ZHCs and these self-reported health outcomes. We then present estimated conditional associations between ZHCs, OPC forms and self-reported ill health, controlling for observable worker and job characteristics. Next, we examine the extent to which these associations are robust to the possible confounding effects of *unobserved* differences between jobs and workers following the approach of Oster (2019). We also present estimated associations for specific socio-demographic groups, occupations and industries, to examine the extent to which any such associations are concentrated among particular groups of workers or particular parts of the labour market. Finally, we discuss different potential explanations for the estimated associations between ZHC employment and self-reported ill health, from sorting of workers with poor health into ZHC employment to the detrimental effects of ZHC employment on health. We draw on results from instrumental variables estimation, instrumenting an individual's contract type with shares of contract types within industries, to inform this discussion. In all three respects, we go beyond Henderson (2019). We conclude by discussing potential policy implications.

We show that workers employed on ZHCs are more likely to report a long-term health condition than workers employed on other types of contract, with the main driver being that they are almost twice as likely to report mental ill health. These associations survive conditioning on an extensive set of observable individual, job and contextual characteristics, and are robust to sensitivity analysis designed to explore the likely impact of biases due to omitted *unobserved* confounders. The conditional association with mental ill health is also stronger than the equivalent associations for several OPC types, which with one exception are insignificantly different from zero. We then show that estimated associations vary little between most demographic groups, although they appear stronger for younger workers and for UK-born workers than for foreign-born workers, and there is tentative evidence that they may be disproportionately concentrated in parts of the economy where underlying job instability is likely to be higher. We argue that both worker sorting and detrimental health effects of ZHCs are likely to contribute to these estimated associations, with implications both for targeting of support and regulation of ZHC employment.

2 | ZHCS IN THE UK

Although there is no universally accepted single definition of a ZHC, even within the UK (Adams & 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 & Williams, 2019), although recent evidence suggests they are no lower paid than equivalent non-ZHC jobs (Farina et al., 2021). 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 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, for example 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).

3 | METHODS

The data used in this article are drawn from the UK QLFS. We pool together data over the period 2015–2018 – pre covid-19 – and restrict analysis to individuals aged 16+ years in employment, excluding the self-employed.[1] The QLFS 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, QLFS 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, annualized hours contract, term time work, job-sharing, 9-day fortnight, four-and-a-half day week, ZHCs, 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 per cent in 2006 to 2.8 per cent (or 901,000 workers) in 2017, approximately where it remained in 2019, the final year before the covid-19 pandemic.

Our focus on the period 2015–2018 is also 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 QLFS 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.[2] Even if we can reasonably rule out such *systematic* under-reporting of ZHCs in the QLFS from 2015 onwards, however, we cannot rule out the possibility that these contractual arrangements continued to be measured with random 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.

QLFS 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 (i.e. answering yes to the above question), the LFS asks respondents to indicate what type or types of health problem they suffer from, drawn from a list of 17 options. Given the 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

TABLE 1 Descriptive statistics by ZHC status.

	Employed not on a ZHC Mean (St. Dev.)	Employed on a ZHC Mean (St. Dev.)
HOURPAY (2017£)	14.59 (9.61)	9.19 (7.42)
HRRATE (2017£)	9.99 (8.88)	8.69 (4.81)
Working Hours	34.017 (11.04)	23.689 (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

(Continues)

TABLE 1 (Continued)

	Employed not on a ZHC Mean (St. Dev.)	Employed on a ZHC Mean (St. Dev.)
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

(Continues)

TABLE 1 (Continued)

	Employed not on a ZHC Mean (St. Dev.)	Employed on a ZHC Mean (St. Dev.)
Region: West Midlands	0.085	0.091
Region: Yorkshire & Humberside	0.079	0.083
N	235,139	6625

Note: 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 to 2018 LFS. Two measures of hourly wages are presented, that is 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-term health problem	0.251	0.255	0.295	191.23***
N	296,239	235,139	6625	
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	6573	

Note: Each entry reports the (weighted) proportions reporting having a long-term 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. ** indicates statistically significant differences at the 1% level.

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. Because our primary interest is the association between ZHCs and *mental* health, and to mitigate the risk of finding statistically significant associations purely by chance in regression analysis with multiple outcomes, we then further aggregate the five categories other than mental into a single catch-all non-mental category, which we label physical health.

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 per cent of all people in employment, and workers not in ZHC employment, report having a long-term health problem. This figure increases to 30 per cent 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 per cent of ZHC workers reporting a long-term mental health condition compared to approximately 3 per cent 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 shown in Table 1, however, there are numerous other observed differences between ZHC workers and those employed under other contract types, on average, many of which are themselves associated with health outcomes, for example age. There are also differences in job characteristics between ZHC jobs and non-ZHC jobs, other than contract type, which may also be associated with health outcomes, for example their sectoral distribution. Such compositional differences may exaggerate or obscure underlying associations between contract type and self-reported health *other things being equal*. In order to compare the health of workers employed under different contract types *net* of such differences, we, therefore, use a multivariate regression approach to estimate *conditional* associations, that is associations between contract type and health outcomes which adjust for observed demographic and job characteristics that potentially confound the associations of interest. Specifically, we estimate linear regressions for each health outcome of the following form[3]:

$$\Pr(Y_i = 1) = \beta \mathrm{ZHC_i} + \mathrm{X}_i'\delta + \tau_q + \varepsilon_i \tag{1}$$

Y_i is the binary health outcome for individual i, for example poor mental health. $\mathrm{ZHC_I}$ takes value 1 for LFS respondents on a ZHC and 0 otherwise. X_i includes a constant, the demographic characteristics in Table 1 (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). Finally, τ_q is a set of quarter/year dummies and ε_i is the error term. With the partial exception of Henderson (2019), which is limited to estimating ZHC and mental health associations for 25-year-olds in England, this article is the first to estimate conditional associations of this kind between ZHC status and multiple health outcomes across all working ages for the UK.

Even after extensive conditioning on observed worker and job characteristics, however, there may still be *unobserved* differences between workers (and jobs) employed under different contractual arrangements that confound estimated associations. For example, those in disadvantaged local areas may be both more likely to suffer ill health and more likely to be employed under a ZHC contract. While we cannot adjust explicitly for this, we can assess the sensitivity of our estimated associations to unobserved confounders following the approach of Oster (2019), which assumes that changes in estimated coefficients and *R*-squared after inclusion of *observed* controls (i.e. after accounting for observed differences between ZHC and other workers) can be informative about the likely confounding role of *unobservable* differences between ZHC workers and those employed under other types of contract. Among the results presented from this Oster-style analysis are bias-corrected regression coefficients assuming proportional selection from observables and unobservables, that is unobservable factors confound estimated associations in the same direction and to the same degree as observed confounders. Bias-corrected coefficients that remain statistically significant and non-trivial in magnitude suggest associations that are likely to be at least qualitatively robust to the presence of unobservable confounders. We also present estimates of the degree of selection on unobservables required to explain away the Equation (1) estimated associations, that is the estimated ratio of selection on unobservables compared to selection on observables which would be required to 'kill' the estimated coefficient. Estimated δs above one are again suggestive of associations that are likely to be at least qualitatively robust to the presence of unobserved confounders. This is the first article to present this kind of sensitivity analysis in the context of ZHC and health associations.

In addition to estimating (1) on the full sample, we re-estimate (1) splitting the sample by age group, gender, and a series of other individual and job characteristics to examine whether associations between ZHC employment and ill health vary along these dimensions. There are two motivations for this. First, by examining where estimated associations between ZHC employment and ill health are strongest, we might potentially help to target interventions, for example mental health support for ZHC workers, where they are most needed. Second, variation across demographic groups and job types in the degree to which ZHC employment is associated with ill health might offer some suggestive insight into the underlying mechanisms behind associations between ZHC employment and (mental) ill health. Again, this article is the first article to do this in this particular context.

Finally, in an attempt to further tease out whether an underlying causal effect of ZHC employment on health might help to explain estimated associations, we estimate an instrumental variables (IV) model, by two-stage least squares (2SLS), as follows. We first simplify (1) by pooling all other (non-ZHC) precarious contract types into a single category, which we call 'other precarious contracts' (OPC), and by focussing only on mental ill health as an outcome. We then allow both ZHC status and OPC status to be endogenous (perhaps due to sorting or unobserved confounders or both), instrumenting each with contemporaneous national-level variation in ZHC and OPC employment shares at the 4-digit sectoral levels. In doing so, we exploit a strong source of plausibly exogenous variation: workers in a given industry are more likely to be on a ZHC (or OPC) if they are hired in a sector in which ZHCs (OPCs) are more widely used. A similar IV approach is adopted by Wooden et al. (2023) in their study of the effects of precarious employment on fertility and by Moscone et al. (2016) in their study of the effects of fixed-term employment on health (albeit within-firm shares rather than within-sector shares in the latter case).

The first stage of the IV model is specified as follows:

$$\Pr(ZHC_i = 1) = \beta_2\, ZHCR_i + \gamma_2 OPCR_i + Z_i'\delta_2 + \theta_q + \epsilon_i \tag{2}$$

$$\Pr(OPC_i = 1) = \beta_3\, ZHCR_i + \gamma_3 OPCR_i + Z_i'\delta_3 + \phi_q + e_i \tag{3}$$

$ZHCR_i$ ($OPCR_i$) denotes the ZHC (OPC) rate (or share) in individual i's 4-digit industry at the national level. Z_i is a vector of demographic and job characteristics equal to X_i in (1) but without the dummies for other contract forms, θ_q and ϕ_q are sets of quarter/year dummies, and ϵ_i and e_i are error terms. The identifying assumption for this IV approach to be valid is that individual i's health is not affected by the proportion of ZHCs (OPCs) in a given industry other than through own ZHC (OPC) status. This assumption is untestable, but Wooden et al. (2023) points to an existing literature using aggregated labour market conditions to instrument for individual employment status. Furthermore, were there other mechanisms through which industry ZHC shares impacted on individual mental health (e.g. indirect effects or sorting), we might expect the resulting IV estimates to be sensitive to the inclusion or exclusion of demographic or job characteristic covariates, which we can and do test. Ultimately, however, given the untestable nature of the identification assumption for this IV approach, we treat these IV estimates as suggestive rather than anything more. But we also encourage readers to consider them together with the main OLS estimates, the Oster-style analysis and the sub-sample estimates *in the round*.

4 | RESULTS

Table 3 presents estimated conditional associations between ZHC employment, reporting a long-term health condition, and the six categories for main type of health condition, all from our model adjusting for worker, job and contextual characteristics.[4] We also report the conditional association between ZHC employment and our outcome pooling all non-mental health categories into a single physical heath measure, and equivalent conditional associations for other precarious employment types. The omitted category for employment type is permanent, that is workers who do not report a ZHC or any kind of temporary contractual arrangement.

The estimates in column (1) demonstrate that ZHC workers are more likely to report a long-term health problem (+2.9 percentage points, on a base of 25 per cent, so approximately 12 per cent higher) compared to permanent workers, even after conditioning on our rich set of observable controls. This association is statistically significant at the 1 per cent 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-term health condition takes the opposite sign but is smaller in magnitude and only marginally statistically significant at standard levels. 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 1.7 percentage points (∼50 per cent) more likely to report suffering from such a condition than other workers, statistically significant at the 1 per cent 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. Note that with seven health outcomes (columns 1–7), there is an increased risk of type-one error, that is one or more associations might appear statistically significant purely at random. However, the associations between ZHC employment and reporting a long-term health problem and ZHC employment and reporting a mental health problem are both sufficiently strong so as to remain statistically significant even after a conservative Bonferroni correction to account for multiple outcomes. There is little evidence that other forms of precarious work are as strongly associated with variations in mental health, with only fixed-term contracts (temporary fixed period) having a small positive association with reporting mental health problems. Only for temporary (fixed period) and temporary (other) do the 95 per cent confidence intervals overlap with that for the ZHC coefficient.

Although the estimated associations between ZHC status and the five other non-mental health conditions are all individually statistically insignificant, they are all positive, and when pooled into a single physical/non-mental category, there is a statistically significant association between ZHC employment and reporting a long-term physical health condition. The estimate shows that

TABLE 3 Regression estimates, ZHCs and health status.

		Main health problem						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Long-term health problem	Muscular	Sensory	Circul./ breath.	Digestive/ kidney/ diabetes	Mental/ depression	Other	All non-mental
ZHC	0.029***	0.004	0.001	0.000	0.001	0.017***	0.005	0.012**
	(0.006)	(0.003)	(0.002)	(0.003)	(0.003)	(0.003)	(0.003)	(0.006)
Perm. Agency Work	−0.011	0.001	−0.002	−0.005	0.000	−0.003	−0.001	−0.007
	(0.007)	(0.004)	(0.002)	(0.004)	(0.003)	(0.003)	(0.003)	(0.007)
Temp.: Agency Work	−0.003	0.004	−0.001	−0.010**	0.001	−0.000	0.004	−0.002
	(0.009)	(0.005)	(0.002)	(0.005)	(0.004)	(0.004)	(0.004)	(0.008)
Temp.: Casual	−0.015*	−0.006	0.002	0.006	−0.005	−0.005	−0.006	−0.009
	(0.009)	(0.005)	(0.003)	(0.005)	(0.003)	(0.004)	(0.004)	(0.008)
Temp.: Seasonal	0.030**	0.003	0.007	0.006	0.011	−0.001	0.005	0.032**
	(0.015)	(0.008)	(0.005)	(0.009)	(0.007)	(0.007)	(0.008)	(0.014)
Temp: Fixed Period	0.023***	0.005	0.005**	−0.000	0.004	0.007**	0.001	0.015***
	(0.006)	(0.003)	(0.002)	(0.003)	(0.003)	(0.003)	(0.003)	(0.006)
Temp.: Other	0.021**	0.014**	0.004	−0.003	0.002	0.003	0.001	0.018*
	(0.010)	(0.006)	(0.003)	(0.005)	(0.004)	(0.005)	(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
R^2	0.051	0.023	0.001	0.019	0.011	0.014	0.009	0.055

Note: Significance at the 10% level is represented by *, at the 5% level by ** and at the 1% level by ***. Entries are estimated coefficients from linear probability models with associated robust standard errors in parentheses. The dependent variables are binary indicators taking value 1 if LFS respondents report long-term 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 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 categorized 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.

ZHC workers are 1.2 percentage points more likely to report such a condition, on a base of 22 per cent, so corresponding to ~5 per cent higher prevalence. Overall, the suggestion is that the positive correlation between reporting a long-term health condition and being employed on a ZHC is driven by both poorer mental health and poorer physical health, with poorer mental health being the primary contributor. There are also some statistically significant associations between OPC forms and individual non-mental health categories scattered over columns (2)–(5) and (7), but the lack of a clear pattern suggests we cannot rule out that at least some of these are randomly rather than meaningfully statistically significant. Again, the pooled outcome for reporting a physical/

TABLE 4 Sensitivity of ZHC regression estimates to selection on unobservables.

	β (from Table 3)	β^* (with proportional selection)	δ^* (degree of selection required for zero association)
Long-term health condition	0.029***	0.025***	6.63***
	(0.006)	(0.006)	(2.06)
Mental health condition	0.017***	0.014***	4.44***
	(0.003)	(0.003)	(1.17)
Physical health condition	0.012**	0.011**	16.45
	(0.006)	(0.005)	(617.74)

Note: Significance at the 10% level is represented by *, at the 5% level by ** and at the 1% level by ***. Entries are estimated coefficients from linear probability models with associated robust standard errors in parentheses. The dependent variables are binary indicators taking value 1 if LFS respondents report a long-term health problem, that their main health problem is mental ill health, or that their main health problem is one of the aggregated non-mental health categories, that is muscular, sensory, circulatory/breathing, digestive or other. 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 categorized 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. In column 2, β is the estimated pollution coefficient from Table 3, that is estimated from Equation (1). In column 3, β^* is the (Oster) bias-adjusted pollution coefficient assuming proportional selection on observables and unobservables (δ equals 1), with bootstrapped standard errors (with 1000 replications). In column 4, δ^* is the degree of selection on unobservables required to generate a pollution coefficient of 0. These (columns 3 and 4) parameters assume a maximum R-squared value of 1.3 times that for Equation (1), following Oster (2019).

non-mental health condition is useful in this respect, suggesting positive associations between poor physical health and seasonal employment, fixed-term employment and ‘other’ precarious employment. In each case, and in contrast to ZHC employment, it is this physical/non-mental association that drives most of the overall association between these contract types and reporting a long-term health condition of any kind.

Table 4 presents summary estimates from the sensitivity analysis examining robustness to the potential additional confounding effects of *unobserved* differences between ZHC workers and jobs and non-ZHC workers and jobs, following the approach of Oster (2019). The first column reproduces the estimated βs for ZHCs from Table 3, for reporting a long-term health condition, a mental health condition and a non-mental/physical health condition. The second column presents estimated βs assuming proportional selection on observable and unobservable confounders. For all three health outcomes, these bias-corrected estimated associations are very close in magnitude (only slightly smaller in each case) to the estimated associations conditioned only on observables, and all remain statistically significantly different from zero at conventional levels. The third column reports estimates of the ratio of selection on unobservables compared to selection on observables which would be required to get a zero association between ZHC employment and the health outcomes. In each case, the ratio is substantially higher than 1, that is unobservables would need to be 6.6, 4.4 or 16.5 times as confounding as observables to explain away the estimated associations between ZHC employment and reporting a long-term health condition, mental ill health and physical ill health, respectively. The suggestion is that the estimated associations between ZHC employment and the health outcomes presented in Table 3 are highly robust to the likely confounding impacts of unobserved differences between ZHC and non-ZHC workers and jobs. Note, however, that the estimates in Table 4 rest on a particular assumption about the ratio of the maximum possible R-squared (were we able to condition on currently unobserved confounders) and the actual R-squared in (1). We follow the suggestion of Oster (2019) in the specific assumption we make in this case (see Table 4 for details) but sensitivity to alternative assumptions cannot be ruled out.

In Table 5, 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 and different job characteristics. Specifically, we split the sample by age, gender, education, tenure, aggregate industry, occupation, citizenship/migrant status and public/private sector. For all three health outcomes, estimated associations are either positive or non-significant for all demographic groups; there is no group for whom the association between ZHC status and reporting a poor health outcome is negative and statistically significant. This is also the case when the sample is split by job characteristics, with the (likely coincidental) exception of the agriculture/fishing sector. But are there potentially meaningful variations in estimated magnitudes and statistical significance across groups? Given the sheer number of estimates presented in Table 5 in what follows, we focus solely on associations between ZHC employment and mental ill health.

Estimated associations between ZHC status and poor mental health are statistically significant and similar in magnitude for 18–24 s, 25—34 s and 35—49 s, but smaller and non-significant for 50+ year olds. This may in part reflect the use of ZHC jobs as part of 'winding-down' retirement strategies and is consistent with LaMontagne et al.'s (2014) finding that the association between mental health and casual work in Australia is different for older workers than for younger workers. There is no difference by gender; both males and females show a statistically significant association between ZHC employment and poor mental health, and of similar magnitude. Contrast this with evidence of gender differences in the association between fixed-term employment and mental health in Australia from Richardson et al. (2012), and temporary employment and mental health in Britain from Robone et al. (2011). We find an inverted U-shaped relationship between ZHC and mental ill health in terms of education level; those at either extreme show no conditional association, while those in the middle education categories show statistically significant associations with broadly similar magnitudes. Again, there is some existing evidence for mixed associations between temporary work and health by education level in Richardson et al. (2012) and Robone et al. (2011). The association between ZHC employment and mental ill health is larger in magnitude (and only statistically significant) for those born in the UK compared to those born outside the UK.

Turning to job characteristics, there is an apparent public/private sector split, where we only see evidence of a statistically significant association between ZHC employment and mental ill health in the private sector not the public sector, which could reflect greater underlying job instability or a greater degree of employer-driven sorting in the private sector than in the public sector. Furthermore, although small sample sizes are an issue in some cases here, associations between ZHC employment and mental ill health appear to be concentrated in some industries and occupations more than others, although at this level of aggregation, it is unclear to what extent this might reflect differences in underlying job stability and employer-driven variability in hours. In particular, associations appear to be stronger for those in the restaurant/hotel sector and those in the public administration, education and health sector (this includes care work). The occupations that show a statistically significant association between ZHCs and mental ill health are administrative, personal services, process/plant/machine operatives and elementary occupations.

TABLE 5 ZHC regression estimates by worker and job characteristics.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	LONG-TERM HEALTH CONDITION			MAIN CONDITION MENTAL			MAIN CONDITION PHYSICAL		
	β	s.e.	*N*	β	s.e.	*N*	β	s.e.	*N*
PANEL 1: AGE									
16–24	0.026***	0.010	24,267	0.021***	0.006	24,248	0.004	0.009	24,248
25–34	0.042***	0.014	48,099	0.025***	0.009	48,044	0.017	0.012	48,044
35–49	0.039***	0.013	83,284	0.018**	0.007	83,136	0.020	0.013	83,136
50+	0.019	0.013	77,218	0.006	0.005	76,959	0.013	0.013	76,959
PANEL 2: GENDER									
Male	0.023**	0.009	112,687	0.015***	0.004	112,497	0.007	0.009	112,497
Female	0.027***	0.008	120,181	0.017***	0.004	119,890	0.009	0.008	119,890
PANEL 3: EDUCATION									
Degree	0.014	0.014	78,538	0.001	0.006	78,381	0.014	0.013	78,381
Higher Education	0.068***	0.021	23,137	0.022**	0.011	23,085	0.045**	0.020	23,085
GCE A Level	0.013	0.011	52,136	0.016***	0.006	52,025	−0.003	0.010	52,025
GCSE Level	0.044***	0.012	46,800	0.028***	0.007	46,701	0.015	0.011	46,701
Other Education	0.039**	0.019	18,019	0.022**	0.010	17,971	0.016	0.018	17,971
No Education	0.007	0.023	14,238	0.010	0.010	14,224	−0.002	0.022	14,224
PANEL: TENURE									
0–11 months	0.031***	0.010	37,349	0.031***	0.006	37,268	−0.000	0.009	37,268
12+ months	0.027***	0.008	195,519	0.008**	0.003	195,119	0.020***	0.007	195,119
PANEL 5: INDUSTRY									
Agri/Fish	−0.001	0.090	1526	−0.023**	0.010	1523	0.022	0.090	1523
Banking	0.055***	0.020	36,175	0.005	0.008	36,105	0.049**	0.019	36,105
Construction	−0.023	0.043	11,442	0.013	0.018	11,428	−0.035	0.041	11,428
Restaurants/Hotel	0.031***	0.010	44,031	0.024***	0.006	43,953	0.007	0.009	43,953
Energy	0.111	0.095	4446	0.009	0.033	4440	0.102	0.092	4440

(Continues)

BJIR An International Journal of Employment Relations LSE

TABLE 5 (Continued)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	LONG-TERM HEALTH CONDITION			MAIN CONDITION MENTAL			MAIN CONDITION PHYSICAL		
	β	s.e.	N	β	s.e.	N	β	s.e.	N
Manufacturing	0.018	0.027	24,141	0.014	0.012	24,084	0.006	0.025	24,084
Other Services	0.022	0.020	10,631	0.001	0.010	10,611	0.019	0.019	10,611
Publ. Adm., Educ., Health	0.028**	0.011	80,697	0.017***	0.006	80,512	0.011	0.010	80,512
Transport	−0.001	0.026	19,779	0.019	0.012	19,731	−0.020	0.024	19,731
PANEL 6: OCCUPATION									
Managers & Senior Off.	0.064	0.043	23,056	0.032	0.022	23,009	0.033	0.041	23,009
Professional	0.013	0.023	50,397	0.001	0.009	50,291	0.011	0.022	50,291
Associate Professions & Tech.	0.027	0.026	31,614	−0.002	0.011	31,553	0.026	0.025	31,553
Admin. & Secretarial	0.017	0.026	28,185	0.029**	0.014	28,115	−0.012	0.023	28,115
Skilled Trades	0.046*	0.027	17,730	−0.007	0.008	17,701	0.052**	0.026	17,701
Personal Service	0.048***	0.013	22,649	0.024***	0.007	22,599	0.024**	0.012	22,599
Sales & Costumer Service	0.024	0.021	19,532	0.000	0.010	19,489	0.020	0.019	19,489
Process, Plant and Machine Op.	0.012	0.022	14,223	0.028***	0.011	14,202	−0.015	0.021	14,202
Elementary	0.015	0.010	25,482	0.017***	0.006	25,428	0.001	0.010	25,428
PANEL 7: MIGRANT STATUS									
Born in UK	0.033***	0.007	197,459	0.020***	0.004	197,051	0.013**	0.007	197,051
Born outside UK	0.012	0.012	35,409	0.006	0.005	35,336	0.006	0.011	35,336
PANEL 8: PUB. SECT									
Private Sector	0.029***	0.007	168,888	0.019***	0.004	168,558	0.010	0.006	168,558
Public Sector	0.012	0.016	63,980	0.003	0.007	63,829	0.008	0.015	63,829

Note: Significance at the 10% level is represented by *, at the 5% level by ** and at the 1% level by ***. Entries are estimated coefficients and associated robust standard errors and sample sizes from Equation (1) estimated on each subsample. The dependent variables are binary indicators taking value 1 if LFS respondents report a long-term 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 categorized 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.

Again, although there is a risk of over-interpretation related to multiple hypothesis testing here, particularly with the industry and occupation splits, the statistically significant associations that do show through are *highly* statistically significant and would survive conservative adjustment. Finally, although statistically significant associations between ZHC employment and reporting a mental health condition are found for those in jobs with both less than and more than 1 year's tenure, the magnitude of the association is larger for those in the first year of ZHC jobs.

Table 6 presents estimates from various specifications of the IV model for mental ill health described earlier in the article. In each case, it is the second stage estimates that are presented in the table, with key first-stage estimates listed in the table notes. Note that in each case the first stages suggest a strong correlation between the instruments and own contract type, that is the instruments are relevant. In other words, you are much more likely to be employed on a ZHC (OPC) if you work in an industry with a high prevalence of ZHCs, and vice versa. The first column presents Ordinary Least Squares (OLS) equivalents to (1) but with other (non-ZHC) precarious contract types merged into the catchall OPC category. Note that, on average, there is no association between being employed on an OPC contract and mental ill health. The next column presents the standard IV estimates. Again, there is a clear and statistically significant ZHC coefficient but a zero for OPC. If our identifying assumption holds, then the only mechanism through which this estimated effect can operate is via a causal effect of ZHC employment on mental health. Under the identifying assumption, this estimate provides evidence of a detrimental impact of ZHC employment on mental health. Because we cannot establish the validity of the identifying assumption beyond a reasonable doubt, however, we treat this evidence as being merely suggestive of the existence of such a causal effect. Having said that, the IV estimate of the ZHC effect on mental health is qualitatively robust to several sensitivity analyses, which together lend further support to this interpretation. Specifically, in the remaining columns, we present IV estimates from versions of the model with no demographic and job characteristic covariates, where the ZHC and OPC shares vary at the region-sector level rather than the national-sector level, and where we instrument own ZHC/OPC status with historical rather than contemporaneous sectoral ZHC/OPC shares.

5 | DISCUSSION

This article shows that workers employed under ZHCs disproportionately suffer poor health, in particular, poor mental health. We provide the first quantification of this association for the UK using nationally representative QLFS data, showing that even after conditioning on observed worker and job characteristics, ZHC workers are approximately 50 per cent more likely to report mental ill health than observationally equivalent workers in permanent jobs, on average. We contrast the association between ZHC employment and mental ill health with those between other precarious employment types and mental ill health, demonstrating that the association is stronger for ZHC employment than for any other form of employment identifiable in the data. We also show how this association varies between ZHC workers with different observed characteristics and between ZHC jobs in different sectors and occupations. Finally, we use Oster-style analysis to show that it is highly unlikely that the estimated association between ZHC employment and mental ill health presented here is driven by remaining, *unobserved* confounders. Our IV estimates are also consistent with this interpretation.

It, therefore, seems likely that most of the estimated conditional association between ZHC employment and mental ill health presented here is driven either by a causal effect of ZHC employment on mental ill health, by sorting of workers with existing mental health conditions

TABLE 6 IV estimates, ZHCs and mental health.

	OLS	IV (sector shares)	IV (sector shares, no controls)	IV (sector/region shares)	IV (historical sector shares)
ZHC	0.016***	0.030**	0.071***	0.020**	0.074***
	(0.003)	(0.014)	(0.012)	(0.008)	(0.027)
OPC	0.002	0.010	0.023***	0.008*	−0.023
	(0.002)	(0.009)	(0.008)	(0.005)	(0.015)
Demo. Characteristics	Yes	Yes	No	Yes	Yes
Job Characteristics	Yes	Yes	No	Yes	Yes
Regional Dummies	Yes	Yes	No	No	Yes
Quarter Dummies	Yes	Yes	No	Yes	Yes
N	232,387	232,387	232,387	232,381	232,360
R^2	0.014	0.014	−0.002	0.014	0.011
Kleibergen–Paap rk Wald F		1521.71***	1696.90***	3218.64***	409.86***

Note: Significance at the 10% level is represented by *, at the 5% level by ** and at the 1% level by ***. Entries are estimated coefficients from OLS (column 2) and 2SLS estimation (columns 3–6) with associated robust standard errors in parentheses. The dependent variable is a binary indicator taking value 1 if the LFS respondent reports a long-term mental health problem. ZHC and OPC are binary indicators denoting employment in a ZHC job or other precarious contract job (permanent or temporary agency work, casual, seasonal, fixed period or other temporary contract), respectively. 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 categorized 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. In the first stage, ZHC and OPC status is (separately) regressed on the national share of ZHCs and OPCs in overall employment in the individual's 4-digit industry in the contemporaneous quarter (columns 3 and 4). The model for column 5 is identical but ZHC and OPC shares are defined at the 4-digit sector level within region (with region dummies, therefore, omitted). The model for column 6 is identical but ZHC and OPC shares are defined as the average shares across Q2 and Q4 of 2009 and 2010. Reading from left to right, starting with column 3, the relevant first-stage coefficients (standard errors) for ZHC share are as follows: 0.970*** (0.022), 1.208*** (0.023), 1.008*** (0.015) and 2.164*** (0.086). The equivalent coefficients (standard errors) for OPC share are: 0.876*** (0.016), 0.964*** (0.016), 0.897*** (0.008) and 0.635*** (0.018). In each of the IV models, we reject the null hypothesis of weak instruments at the 1% level.

into ZHC employment, or some combination of the two. First, consider potential mechanisms for a causal relationship running from ZHC employment to mental ill health, that is a detrimental mental health effect of ZHC employment. Existing evidence suggests that job insecurity can impact adversely on psychological stress (Cheng & Chan, 2008; Ferrie, 2001; Green, 2015; Sverke et al., 2002). This mechanism is likely to be relatively strong for ZHCs, given their extreme precarity, and may be exacerbated by working-hours fluctuations over which the worker has limited control. Other causal mechanisms for detrimental effects on mental ill health and ill health more generally might include higher-levels of sickness-related presenteeism among ZHC workers given such jobs typically offer limited entitlement to paid sick leave (Skagen & Collins, 2016) and higher risk of work-related injuries and illness given precarious jobs may offer 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). The existence of such a causal effect is supported by qualitative studies showing a perceived detrimental impact of ZHC employment on health, particularly mental health (e.g. Ball et al., 2017; Ndzi et al., 2017), echoed by anecdotal evidence in media reports (e.g. O'Connor, 2019).

Second, consider worker sorting.[5] One potential mechanism for this is that workers with mental ill health may sort themselves into ZHC employment (self-sorting) because ZHCs offer (or are perceived to offer) greater flexibility from the worker's perspective in terms of hours and shifts than other employment contracts, which may help in the management of long-term health conditions. 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 seek ZHC work. Employers may also sort workers with mental ill health into ZHC jobs if workers with mental health conditions are disproportionately excluded from non-ZHC jobs, as might be expected in a segmented labour market. Employers may also disproportionately use ZHC jobs as a screening mechanism for workers with mental ill health. To the extent that ZHC jobs offer workers genuine two-sided flexibility, employers may also use them in order to better accommodate additional needs for flexibility among workers with mental ill health. These potential mechanisms can also help to explain associations between ZHC employment and other health conditions.

With cross-sectional data, we are unable to quantify the relative importance of these potential underlying mechanisms in explaining the conditional associations between ill health and ZHC employment presented here. Nevertheless, the IV analysis presented here potentially offers some insight. Taken at face value, these IV estimates provide evidence of a detrimental causal effect of ZHC employment on mental health, and one of non-trivial magnitude. However, despite their robustness to inclusion/exclusion of covariates, it is difficult to entirely rule out the possibility that some sorting might remain even in the IV estimates, whereby those with better/worse mental health sort into industries with higher/lower ZHC shares. The IV estimates are, therefore, best interpreted as being *suggestive* of a causal effect.

There are some suggestive indications in our estimates by worker characteristics and by sector that employer sorting and detrimental health effects of ZHCs may be more salient than self-sorting. For example, the public/private sector contrast suggests self-sorting is unlikely to be the driving force because it is not clear why such sorting would affect the private but not the public sector. In contrast, it seems entirely possible that private sector employers might sort workers with existing ill health into ZHC jobs to a greater extent than public sector employers. Similarly, the public/private sector contrast could be suggestive of at least part of the overall association reflecting a causal effect of ZHC employment on mental ill health if private sector ZHC jobs are more insecure or feature more one-sided flexibility than public sector ZHC jobs. Similar arguments apply to why we find stronger associations between ZHCs and mental ill health in some industrial sectors than others.

The age contrast presented here may also be suggestive of a detrimental ZHC impact on mental health, for example if older workers are less dependent on income (and predictability of income) from ZHC jobs than younger workers. In contrast, it is unclear why younger workers with existing health conditions would self-select into ZHC employment to a greater extent than older workers, although it seems possible that employers might use ZHC contracts to screen younger workers to a greater extent than older workers who are more likely to have an existing employment record. On the other hand, the lack of an association among migrant workers seems more likely to reflect employer sorting of migrant workers into ZHC employment regardless of health in contrast to sorting by health status within the native workforce, rather than a heterogeneous causal effect of ZHC employment.

At first glance, heterogeneous associations by tenure might offer a way of distinguishing between employer sorting and a detrimental causal effect of ZHC employment on mental health. Specifically, if the association between ZHC employment and mental ill health was explained primarily by detrimental health effects of ZHCs, then we might expect a dose–response, that is stronger effects at higher tenures. If employer sorting were the main driver, we might expect the opposite, which is what we observe in practice. However, we must be careful not to over-interpret such differences. Longer-tenure ZHC jobs suggest more stability in the employment relationship, removing an important mechanism for a causal effect at longer tenures but not shorter tenures. Similarly, those with mental ill health may be more likely to get stuck in ZHC jobs over the longer term, so we might expect stronger sorting effects at longer tenures than shorter tenures.

The relative importance of worker sorting and detrimental health effects of precarious employment in explaining associations between ill health and precarious employment also remains an open question in the wider literature. The most promising strand of the literature in this respect exploits longitudinal data to estimate the extent to which such associations might be driven by an underlying casual effect of precarious employment on ill health. As yet, however, this literature has presented little conclusive evidence for such an effect, at least on average. For the UK, for example, Bardasi and Francesconi (2004) and Robone et al. (2011) find no effect of temporary employment on mental or general health. Similarly, for Australia, Richardson et al. (2012), LaMontagne et al. (2014) and Hahn et al. (2021) find no evidence of an adverse effect of casual employment on mental or other measures of health. We should be wary of extrapolating these conclusions to ZHCs in the UK context, however, given the particularly precarious nature of ZHCs and given we show here that the association between mental health and ZHC employment is stronger than that for other contractual forms.

Even in the absence of clear evidence on whether ZHC employment causes detrimental health effects, the strong association between ZHC employment and mental health presented here can potentially help to inform worker and firm behaviour, along with policy design. For example, improving understanding about how worker health varies across contract types might help firms and policy makers to target appropriate support, and might help workers make more informed decisions about whether and where to accept ZHC jobs. Furthermore, given the suggestive evidence presented here that worker self-sorting does not explain the association between ZHCs and poor mental health, policy makers considering further regulatory restrictions or even outright bans of ZHCs can likely discount any concerns that such steps might restrict the employment opportunities of workers with existing mental health conditions. Ultimately, however, how we respond to the growth in ZHCs, and the likely impacts of tightening regulatory restrictions or banning ZHCs, needs to be informed by clear evidence on the causal impact of ZHCs on worker wellbeing, including on workers' health. Only when we have such evidence can we make informed predictions about the extent to which regulating or restricting the use of ZHCs might impact positively on worker wellbeing, including on mental health.

The fact we are not able to unambiguously identify a causal effect of ZHC employment on mental health is the main shortcoming of the current study. Doing so is difficult with cross-sectional data, and the kind of IV approach adopted here and in an earlier article by Moscone et al. (2016) for temporary employment, rests on assumptions concerning instrument exogeneity that are essentially untestable. A promising avenue for further research is likely to be exploiting longitudinal data along the lines of the recent strand of the literature that does so for other forms of precarious employment. It is only now, however, that such an approach is becoming possible for the UK, with the recent inclusion of a question on ZHC employment in the UK Household Longitudinal Study from wave 8 (2016/17), albeit only in every other wave. Further research might also examine the extent to which the relationships between ZHC employment and worker health may have strengthened during the covid-19 pandemic given the concentration of many ZHC jobs in key service industries that likely exposed ZHC workers to higher risk of covid-19 than otherwise comparable workers.

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DATA AVAILABILITY STATEMENT

The data used in the manuscript are all publicly available via the UK Data Services portal. Codes reproducing the reported results may be obtained upon request via the author.

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ENDNOTES

[1] Including the self-employed in the analysis sample, with a self-employed dummy variable included in the set of precarious employment types, does not affect our conclusions.

[2] A further motivation for this restriction is that prior to 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).

[3] Conclusions are robust to non-linear (probit) estimation.

[4] The full set of results for the controls reveal a range of commonly found associations. For instance, there is strong age gradient in both long-term health and mental health problems. While, women have a lower probability of long-term health but higher probability of mental health problems.

[5] Although we are aware of no evidence on this specifically for ZHC employment, Dawson et al. (2015) use 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.

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