



Working Paper 4/2025
Deliverable 3.3

**Home Advantage or Hidden Strain?: The Mental Health
Effects of Working from Home Across Gender, Childcare
Status, and Occupational Class Before and Since the
Pandemic**

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Home Advantage or Hidden Strain?

The Mental Health Effects of Working from Home Across Gender, Childcare Status, and Occupational Class Before and Since the Pandemic

Abstract

Although working from home (WFH) is promoted as a policy supporting work-life balance, whether it benefits mental health remains unclear. Few studies have examined how these effects vary across social groups and addressed selection issues complicating causal inference. We use two-way fixed effects models to analyse changes in GHQ-12 mental health scores among 39,863 participants in the UK Household Longitudinal Study (2009–2023). We reduce selection bias by using an occupation-level WFH measure derived from the UK Labour Force Survey. Before March 2020, increased WFH in men's occupations improved their mental health. For women, it benefited those in routine jobs but worsened outcomes for professionals. The pattern reversed from March 2020, with WFH positively impacting the mental health of professional women but not that of men or women in routine jobs. These findings highlight the importance of social positions and institutional contexts in shaping the mental health effects of WFH.

Sociology of stress scholars have conceptualised mental health inequalities as the result of unequal exposure and vulnerability to chronic stressors across axes of social inequalities such as age, gender, class, and race (Fenwick and Tausig 1994; Pearlin 1999; McLeod 2015). Work plays a central role in this process as it concentrates multiple sources of strain, including job insecurity, high job demands, and blurred work–life boundaries (Clougherty et al. 2010). However, the nature of work has undergone profound transformations in recent decades, driven by work intensification, the rise of digital technologies, and increasing job insecurity (Kelly and Moen 2020). Therefore, examining how new work practices emerging from these shifts affect different groups of workers is essential for understanding the dynamics of mental health inequalities (Kelly et al. 2011; Fan and Moen 2023).

One major shift is the rise of working from home (WFH), which accelerated during the COVID-19 pandemic and has now become a central feature of many jobs (Barrero et al. 2023). By reducing commuting time (Barrero et al. 2020) and enhancing job autonomy and schedule control (Gajendran and Harrison 2007), WFH expands individuals' temporal and psychosocial resources. As such, it may foster workers' mental health by alleviating work-related stress and conflicts between work and family demands (Voydanoff 2005). The benefits may be particularly pronounced for women, who are disproportionately affected by such conflicts. Not

only does WFH help them remain employed during periods of increased caregiving demands (Chung and Van der Horst 2018), but it may also reduce the strain they face when employed. Similarly, WFH may enhance fathers' mental health by allowing them to take on more childcare (Carlson et al. 2021), which is associated with higher emotional rewards for them (McDonnell et al. 2019).

However, it remains unclear whether WFH uniformly supports or undermines workers' mental health. Studies show that WFH can increase mothers' involvement in childcare and housework, reinforcing the gendered division of unpaid labour (Chung and Booker 2023) and, in turn, worsening their work–family conflicts (Yucel and Chung 2023). Consequently, women with caregiving responsibilities may remain employed yet experience chronic strain and higher mental distress when WFH (Culatti 2014; Beaufils et al. 2023). Similarly, WFH can lead men to work longer hours as work–family boundaries blur, especially in contexts where biases against flexible workers persist (Lott and Chung 2016), thereby potentially harming their mental health (Berniell and Bietendeck 2020). Furthermore, how WFH affects mental health likely varies with workers' caregiving responsibilities and occupational position. Indeed, parenthood increases unpaid labour demands (Craig and Mullan 2010), making them more exposed to work-family conflicts that WFH may either exacerbate or alleviate. Workers' position in the occupational hierarchy influences the degree of autonomy that WFH provides (Clawson and Gerstel 2016), as well as access to alternative resources for managing competing work and family demands, such as flexible schedules, paid leave, and outsourcing (Kim 2020).

In this article, we examine the impact of WFH on mental health across subgroups defined by gender, childcare responsibilities, and occupational class. We use data from the UK Household Longitudinal Survey and the UK Labour Force Survey. The UK provides a relevant context, as it stands out among European countries for the scale and institutionalisation of WFH following

the pandemic: as of 2024, over a quarter of workers reported mainly WFH, and nearly half reported doing so occasionally (Chung and Yuan 2025).

Building on the existing literature, we ask the following questions:

- What is the causal impact of WFH on the mental health of men and women in the UK?
- Do the mental health effects of WFH vary by gender, childcare status and occupational class?
- How have these effects changed before and since the pandemic, which triggered a sharp increase in WFH?

We make several contributions. First, we employ a novel empirical strategy, inspired by Belloni et al. (2022), to estimate the causal impact of WFH on mental health. This strategy addresses selection into WFH, where mental health influences the decision to WFH or the ability to do so. By linking the UK Labour Force Survey to the UK Household Longitudinal Study, we construct a detailed occupational-level measure of WFH prevalence. This measure strongly predicts individual WFH behaviour yet is unlikely to be influenced by individuals' mental health, thereby helping to address reverse causality. Additionally, our models include individual fixed effects, allowing us to examine within-person changes in mental health associated with WFH while controlling for all time-invariant individual characteristics that could confound this relationship. Unlike group-randomised workplace interventions (Moen et al., 2011), our approach enables us to generalise causal claims across diverse social groups, providing a broad picture of the effects of WFH in the UK.

Second, we adopt an intersectional approach (Collins and Bilge 2020) to investigate whether the effects of WFH on mental health vary according to individuals' social structural locations — specifically gender, childcare status, and occupational class. While previous causal studies have primarily focused on gender differences, these effects are likely to be shaped by multiple,

intersecting axes of structural inequalities. Building on the job demands-resources model (Bakker and Demerouti 2007), role strain theory (Greenhaus and Beutell 1985) and the stress process model (Pearlin 1999), we argue that WFH affects workers' mental health through its interaction with job-related stressors and their interference with family demands. We hypothesise that gender, childcare responsibilities and occupational class shape how exposed and vulnerable individuals are to these stressors, and that the mental health impact of WFH is likely to vary accordingly.

Third, we compare the impact of WFH on mental health before and after onset of the COVID-19 pandemic. While numerous studies have examined this relationship either before or during the pandemic (Lunde et al. 2022; Vacchiano et al. 2024), to our knowledge, none have used the same data and analytical approach for both periods. This comparison enables us to question whether broader institutional and normative contexts influence the mental health consequences of WFH. As WFH has become normalised and organisations have adapted during and post-pandemic, it may have become less likely to exacerbate job-related stressors and work–family conflicts, potentially supporting workers' mental health (Fan and Moen 2023a; Fan and Moen 2023b).

1. Background

a. A review of the mechanisms linking working from home and mental health

WFH may benefit the mental health of both men and women by limiting their exposure and vulnerability to stressors stemming from job demands and their interference with family demands. The Job Demands–Resources model (Bakker & Demerouti, 2007) conceptualises work as involving demands (e.g. workload, physical strain, role conflict) that, if not offset by job resources (e.g. autonomy, social support), can lead to job strain. These resources help buffer the chronic strain produced by such demands, thereby protecting mental health. By increasing

job autonomy, WFH provides psychosocial resources that help workers cope with high job demands and reduce job strain (Gajendran and Harrison 2007). It also enhances schedule control and eliminates commuting time (Barrero et al. 2020), making it a ‘boundary-spanning resource’ that helps workers manage competing demands across different life domains (Voydanoff 2005). In doing so, it may reduce work/family conflicts (Laß and Wooden 2023; Baum and Rau 2024), defined as situations where work interferes with family responsibilities, which are known to cause chronic strain that negatively affects mental health in both the short and long term (Greenhaus 2006; Tsukerman et al. 2020). Additionally, WFH availability may strengthen workers’ perceptions of support from managers and organisations, which plays a role in reducing work–family conflicts (Allen 2001).

Conversely, WFH may negatively affect workers’ mental health by increasing, rather than decreasing, job strain and work-family conflicts. Technologies enabling WFH foster a ‘norm of responsiveness’ that raises job demands through a faster work pace and a culture of ‘always-on’ (Chesley 2014; Mazmanian et al. 2013). At the same time, employees who WFH often report lacking supervisor and colleague support, which is a critical resource for coping with high demands (Ha 2021). This imbalance between rising demands and diminishing resources may heighten job strain and offset the autonomy benefits of WFH (Bakker and Demerouti 2007). Moreover, as WFH blurs the physical and psychosocial boundaries between work and family (Clark 2000), it may increase rather than decrease conflicts between the two spheres, acting as ‘boundary-spanning demands’ (Voydanoff 2005). Evidence indeed suggests that workers who WFH frequently multitask between professional and caregiving duties (Andrew et al. 2020) and struggle to disconnect from work (Mazmanian et al. 2013; Kelly and Moen 2020). They tend to extend their working hours, especially in the case of men (Glass and Noonan 2016; Chung 2022), or their childcare and housework hours, especially in the case of women (Chung and Booker 2023). These factors together create conditions for increased job

strain, as well as heightened work-to-family and family-to-work conflicts (Yücel and Chung 2023). This may be exacerbated in contexts where the 'ideal worker' norm prevails, whereby employees are expected to work long hours and prioritise work over family responsibilities (Acker 1990; van der Lippe and Lippényi 2020).

b. Existing evidence on the mental health impact of working from home

A large body of research has examined the mental health impact of WFH, with an almost even split between studies reporting negative and positive outcomes (Lunde et al. 2022; Vacchiano et al. 2024). However, few of these studies have used designs that address reverse causality. For example, recent research using UKHLS data shows that when individuals gained access to and started using WFH, their mental health improved, for both men and women (Li and Wang 2022). However, this may partly reflect that healthier individuals are more likely to request or receive WFH due to stronger networks or better relations with employers (Hoque and Bacon 2021). Unobserved factors associated with both WFH access and mental health (such as holding higher-prestige jobs with greater autonomy) may also explain why WFH workers report better mental health (Wheatley 2017). Conversely, studies finding a negative impact of WFH on mental health may be because individuals in poorer health are more likely to request WFH to manage the challenges they face in traditional work environments (Schur et al. 2020).

Among the studies that have used empirical designs to address these issues, a few randomised controlled trials conducted before the pandemic show that employees' control over their workplace improves their mental well-being. For instance, Moen et al. (2016) demonstrated that, in a large tech company, an intervention that increased employees' control over their work location significantly reduced burnout, perceived stress, and psychological distress. Similarly, Bloom et al. (2015) reported higher productivity and job satisfaction among employees who were WFH for nine months in a large Chinese call centre. Experiments in Italy and Belgium

have also found that partial remote working, one to two days per week, improves wellbeing (Angelici and Profeta 2024; Delanoeije and Verbruggen 2020). A recent instrumental variable study based on access to mobile work equipment showed positive mental health effects of WFH, especially for men and middle-aged workers (Denzer and Grunau 2023). However, since these studies use non-representative samples, their findings may reflect firm-specific factors such as organisational culture and how WFH policies are implemented.

c. Heterogeneity by gender, childcare status and occupational class

The mental health impact of WFH likely differs by gender and childcare status. Since women continue to bear a disproportionate share of unpaid labour, they experience greater strain in balancing work and family demands (Sullivan 2021). This strain is more pronounced among mothers with dependent children, whose circumstances increase both childcare and housework (ONS 2019). For these reasons, in this paper we define childcare status as the presence of children under 15 in the household, as this entails daily caregiving demands. Hence, where WFH serves as a 'boundary-spanning resource', facilitating work-family integration, women – especially mothers – may benefit more than other groups. Conversely, where WFH acts as 'boundary-spanning demands', intensifying work-family conflict, their mental health would be more negatively affected (Voydanoff 2005). The latter scenario is likely in organisations where the 'ideal worker' norm and the flexibility stigma dominate (van der Lippe and Lippenyi 2020). Indeed, mothers are particularly vulnerable to this stigma, which compounds existing stereotypes about their professional commitment (Blair-Loy, 2009; Ridgeway and Correll 2004). In response, they may adopt compensatory behaviours, such as increasing their working hours or staying constantly available online (Howcroft et al. 2024), which can heighten work–family conflicts.

By contrast, fathers who WFH tend to increase their working hours, which may negatively affect their mental health (Glass and Noonan 2016; Kim et al. 2020). Meanwhile, their unpaid care time increases only moderately (Carlson et al. 2021), to a level unlikely to generate work–family conflicts and potentially beneficial to their mental health, as evidence shows that fathers’ involvement in childcare relate to emotional rewards (McDonnell et al. 2019). This combination—longer working hours but moderate domestic involvement—reflects the influence of prior household bargaining dynamics, which continue to shape the division of labour under new professional arrangements, enabling men to maintain clear work–family boundaries and expand professional duties rather than take on additional unpaid tasks (Carlson et al., 2021). As a result, we expect fathers’ mental health to remain stable or even improve under WFH, as the potential negative effects of longer working hours may be offset by the benefits of increased job autonomy and engagement in childcare.

The mental health impact of WFH also likely varies by individuals’ occupational class. In this paper, we measure occupational class using the National Statistics Socio-economic Classification (NS-SEC), which captures key differences in income and working conditions (Goldthorpe 2017). We assume these factors to shape both the experience of WFH and its impact on mental health.

Among higher-status occupations, the autonomy and schedule control afforded by WFH may be crucial resources for coping with the 'stress of higher status', which involves greater job demands, longer hours, and higher work–family conflicts (Schieman et al. 2006; McGinnity and Calvert 2009). In this context, WFH may be particularly beneficial to their mental health. However, this may be less true for mothers in these roles, who are more likely to engage in the aforementioned compensatory behaviours due to the significant stigma surrounding flexibility in high-status professional settings (Stone and Hernandez 2013).

In contrast, workers in lower-status occupations generally have less control over their work, with low job and schedule autonomy (Tausig et al. 2011; Damaske et al. 2016). For them, the autonomy brought by WFH may then be a key resource to mitigate stress levels (Kim et al. 2020). While traditional gender norms might suggest greater unpaid care burdens and stronger work-family conflicts for women in these roles when WFH, recent evidence reveals a more complex picture: WFH fosters more egalitarian arrangements in low-income families, with men in lower-status jobs taking on more routine domestic tasks when WFH (Chung and Booker 2023). Given that these groups have fewer resources to outsource caregiving (Clawson and Gerstel 2014), WFH may serve as a critical buffer, enabling households to meet caregiving demands and supporting mental health for both partners.

d. Comparing pre-pandemic and post-pandemic periods

The COVID-19 pandemic has fuelled some research using population-based cohort data and causal identification strategies to estimate the impact of WFH on mental health. Most studies report that WFH negatively affected mental health during the pandemic, especially among women and parents of dependent children (Bertoni et al. 2025, Guéguen and Senik 2023, Senik et al. 2024). However, these findings may not generalise to contexts outside of the pandemic, which introduced specific stressors that partly confounded the mental health risks of WFH (e.g. fear of infection, social isolation, increased caregiving responsibilities) (Perelman et al. 2021). WFH may also have generated specific forms of work–family conflicts and job strain during the pandemic due to the conditions under which it was implemented. In many cases, it was mandatory and full-time (Henke et al. 2016), and workers lacked adequate support, such as sufficient technical infrastructure, IT assistance, or access to a dedicated workspace (Wang et

al. 2021). Caregiving demands likely exacerbated these challenges, particularly following the closure of schools and childcare facilities (Bernhardt et al. 2023).

However, we expect changes as lockdown restrictions in the UK eased from mid-2021. Childcare facilities and schools reopened, and organisations invested in infrastructure and introduced managerial practices supporting WFH (Wheatley et al. 2021). These changes may have helped workers adapt to WFH and benefit from the resources it offers for integrating work and family demands. In parallel, the rise in WFH during the pandemic may have contributed to changing workplace norms, reducing the career penalties and stigma previously associated with flexible working (Schieman et al. 2022; Fan and Moen 2023). This may have also reduced homeworkers' tendencies to work longer hours to reciprocate for the 'gift' of WFH, as observed pre-pandemic (Kelliher and Anderson 2010). As a result, the potential benefits of WFH for mental health may have increased in the context of the pandemic and its aftermath. This idea is corroborated by recent research showing that continuing to WFH is linked to greater reductions in psychological distress compared to returning to the office between October 2020 and April 2022 in the US (Fan and Moen 2023a; 2023b).

WFH may have also reduced its mental health benefits for some. As WFH spread to a broader range of jobs, including those with lower levels of control, it became more closely supervised (Fana et al., 2022), reducing the autonomy that supports mental health. For men, this may have been compounded by increased role conflicts during the pandemic: greater involvement in childcare and housework led some to experience family-to-work conflicts, as they were less accustomed to the daily demands of childcare (Reimann et al. 2020).

2. Method

a. Data: UKHLS and UK LFS

We used data from waves 1 to 14 (2009–2023) of the UK Household Longitudinal Study (UKHLS), also known as Understanding Society. This is a nationally representative panel survey comprising a stratified, clustered sample of around 60,000 individuals from 40,000 UK households. All adults aged 16 and over in these households are interviewed annually. Response rates have remained high, with 63.6% of eligible adults completing a full interview in Wave 14 and an overall attrition rate of approximately 21%.

For our analysis, we restricted our sample to respondents aged 18–65, excluding those who were unemployed, economically inactive, or self-employed. After removing observations with missing data (approximately 10%), the final sample comprised 39,863 respondents, amounting to 169,309 person-wave observations. The health and sociodemographic characteristics of the sample are detailed in Appendix A.

Additionally, we used data from the UK Labour Force Survey (LFS), a continuous, nationally representative survey that samples approximately 40,000 households (100,000 individuals) every quarter. The LFS provides harmonised, high-frequency data on labour market status, hours worked, job characteristics (including WFH) and occupation and industry codes based on SOC and SIC classifications.

b. Measures

i. Mental health outcome: GHQ-12 (UKHLS)

We measured mental health using the 12-item General Health Questionnaire (GHQ-12), a widely used and validated screening tool designed to detect psychological distress and evaluate overall mental functioning (Goldberg et al., 1997). The GHQ-12 collects self-reported information on various symptoms, such as loss of concentration, sleep disturbances, difficulty making decisions, unhappiness and loss of confidence. Respondents rate each item on a four-point Likert scale ranging from 0 (better than usual) to 3 (much less than usual). We computed

a summary index ranging from 0 to 36, with higher scores indicating greater psychological distress.

In sensitivity analyses, we also assessed the robustness of our findings across mental health indicators using the Mental Component Summary from the 12-item Short Form Health Survey (MCS-12), a validated measure of psychological well-being (Huo et al., 2018). Scores are standardised 0–100, with higher values indicating better mental health.

ii. Working from home (UK LFS)

In the UKHLS, respondents were asked every two waves whether their employer offered the option to WFH and whether they used it. This self-reported WFH measure has been used to study the health effects of flexible working (Chandola et al. 2019, Li and Wang 2022), but as discussed above, introduces selection issues.

To address this issue, we used an external, occupational-level measure of WFH prevalence that we built from the UK LFS. In the LFS, respondents are asked: "In your main job, do you work mainly... 1) in your own home, 2) in the same grounds or buildings as your home, 3) in different places using home as a base, or 4) somewhere quite separate from home?". We classified responses 1 to 3 as indicating as "mainly working from home" and computed, for each quarter and each 4-digit SOC2010 occupation, the proportion of individuals mainly WFH.

This occupation- and quarter-specific prevalence was then assigned to UKHLS respondents based on their 4-digit SOC2010 occupation and the quarter in which they were interviewed. As a result, each UKHLS respondent was matched with a measure of the prevalence of WFH in their occupation at each wave, as independently observed in the LFS.

This variable, which we refer to as average occupational-level WFH, reflects typical WFH use in the respondent's occupation at the time of interview. As individuals' mental health is unlikely

to affect WFH prevalence at the occupational level, this measure helps reduce selection bias and allows us to approximate a quasi-experimental design by leveraging exogenous variation in WFH. It also offers more granular, wave-by-wave tracking of WFH compared to the biennial UKHLS self-reports. Furthermore, our occupation-level measure is based on the proportion of individuals who report ‘mainly’ WFH, capturing regular and intensive WFH use. The UKHLS question asks only whether respondents use WFH at all and may therefore include occasional or limited use.

iii. Childcare status and occupational class (UKHLS)

To describe childcare responsibilities, we created a variable indicating whether individuals live with at least one child under the age of 15. This cutoff reflects children primarily in the pre-secondary and early secondary school period, when parental time demands are higher due to children’s limited autonomy (ONS, 2019). It therefore represents substantial childcare responsibilities while maintaining a sufficient sample size. Alternative specifications were tested (e.g. presence of children aged under 3 or 11 in the household; number of co-resident children), yielding similar results but with less precise estimates.

Occupational class was measured using the three-category NS-SEC version, which groups occupations by the difficulty of supervising work and the specificity of skills required (Goldthorpe 2017). The first category, managerial and professional occupations, includes managers, senior professionals, and higher-level administrative roles (e.g., company directors, lawyers, doctors, professors), characterised by high skill-specificity and control over work. The second, intermediate occupations, covers clerical, sales, and associate professional roles (e.g., office administrators, nurses, teaching assistants), with moderate autonomy, responsibility, and semi-routine tasks. The third, routine and manual occupations, comprises elementary service

and manual roles (e.g., cleaners, security guards, factory operatives, delivery drivers), involving lower autonomy, routine tasks, and higher supervision¹.

c. Modelling strategy

First, we verified that our occupational-level WFH measure is a valid proxy for individual WFH behaviour by comparing its trend from 2010 to 2023 with self-reported data. We also tested this association using two-way fixed-effects models, controlling for the same variables as in the mental health analyses described below.

We then estimated linear two-way fixed-effects regression models to examine the impact of occupational-level WFH on GHQ-12. Individual fixed effects controlled for time-invariant characteristics of individuals, whether observed (e.g., race, education, family background) or unobserved (e.g., attitudes), while time fixed effects accounted for period-specific factors affecting all individuals at a given time (e.g., economic recessions). This modelling strategy increases confidence in our estimates showing causal effects (Allison 2009). Nonetheless, our fixed-effects models cannot fully account for changes over time in who engages in WFH or in the conditions under which WFH is implemented. As a result, estimated effects may partly reflect shifts in the composition of homeworkers or in the characteristics of WFH. We therefore interpret our estimates as capturing the average effects of WFH under the prevailing conditions at each period rather than a stable, exogenous treatment effect.

Building on this framework, our models are specified as:

$$Y_{it} = \beta_1 \cdot WFH_{jt} + \beta_2 \cdot X_{it} + \alpha_i + \lambda_t + \varepsilon_{it}$$

¹ Using the three-category NS-SEC allows to capture meaningful distinctions in work characteristics while maintaining sufficient group sizes and within-group WFH variation. We tested alternative specifications using finer NS-SEC categories, which yielded consistent results, yet with less precise estimates.

where Y_{it} is the mental health score for individual i at time t , WFH_{jt} is the average WFH prevalence in occupation j at time t , X_{it} is a vector of time-varying controls, α_i are individual fixed effects, λ_t are wave fixed effects, and ε_{it} is the error term.

We estimated three sets of models to examine the impact of average occupational-level WFH on mental health and its heterogeneity by gender, childcare responsibilities, and occupational class. First, we included an interaction term between average occupational-level WFH and gender to assess gender differences in the effect. Second, we estimated models stratified by gender that included an interaction between average occupational-level WFH and childcare status. Third, we similarly estimate gender-stratified models interacting average occupational-level WFH with occupational class to explore how this effect varies along the occupational gradient. All analyses were conducted separately for the pre-pandemic (before March 2020) and post-pandemic (from March 2020 onwards) periods.

In all models, we controlled for time-varying socio-demographic and work-related factors likely to influence both occupational changes and mental health outcomes. These include age (mean-centered), age squared, marital status (never married, married/cohabiting, divorced/separated/widowed), household composition (presence of children and age of the youngest child), logged real net monthly income, working hours, and job changes between waves.

For all models, we report both parameter estimates and average marginal effects to facilitate interpretation and comparison across subgroups.

3. Results

a. Average occupational-level WFH in UK LFS and self-reported access to and use of WFH in UKHLS

Figure 1 shows the evolution of average WFH between 2010 and 2023, comparing the measures from the UK LFS and the UKHLS. The proportion of individuals reporting access to and use of WFH in the UKHLS follows closely the same pattern as the proportion of those reporting mainly WFH in the UK LFS, for both men and women. Both sources show a consistent upward trend, with a marked increase from 2020 onwards. As shown in Appendix B, this increase is concentrated in intermediate and professional occupations, while routine occupations show no such trend. Nevertheless, some routine roles (e.g. call and contact centre roles, telephonists, communication operators, market research interviewers, customer service representatives, fitness instructors) have experienced an increase in WFH over time. This variation likely contributes to the findings presented later, on the impact of average occupational-level WFH on the mental health of individuals in routine jobs.

It is important to note that, although WFH became widespread in 2020 due to lockdowns, it was only in the summer of 2021, once the official lockdown restrictions were eased, that workers began reporting home as their main workplace. In other words, despite the widespread adoption of WFH during the pandemic, employees initially regarded it as a temporary measure rather than a permanent shift. It was not until 2022 that workers widely recognised WFH as a lasting work arrangement, explaining the increase in its reported prevalence during that period.

Appendix C presents results from two-way fixed effects models that examine the relationship between average occupational-level WFH and individuals' self-reported access to and use of WFH. They confirm the occupational-level measure as a robust proxy for individual WFH practices, as increases in WFH prevalence within occupations over time correlate with higher self-reports of WFH use in the UKHLS. Notably, these associations are consistent for both men and women and remain significant when analyses are restricted to the pre-pandemic period, suggesting that this relationship is not solely driven by the pandemic context.

[Figure 1 about here]

b. Gender differences in the effect of WFH on mental health

Table 1 presents results from fixed-effects models estimating the effects of average occupational-level WFH on GHQ-12 scores. It examines how changes in the average proportion of WFH within individuals' occupations impact their GHQ scores, and whether this differs between men and women. For both measures, higher scores indicate better mental health.

Before March 2020, increases in average occupational-level WFH led to significant improvements in men's mental health, as shown by the positive main effect on the GHQ. However, the significant negative interaction term indicates that this benefit was smaller for women. As illustrated in Figure 2, the marginal effect of average occupational-level WFH on the GHQ was positive and statistically significant for men but nonsignificant for women.

From March 2020 onward, this gender pattern reversed. Significant positive interaction terms suggest that women gained greater mental health benefits than men as WFH became more common in their occupations. Marginal effects show that increases in average occupational-level WFH led to significant improvements in women's GHQ scores, while effects for men were nonsignificant. Appendix D shows that MCS-12 results follow the same pattern as the GHQ-12, confirming consistency across mental health measures.

[Table 1 about here]

[Figure 2 about here]

c. Variations in the effect of WFH on mental health with childcare status

Table 2 presents estimates from gender-stratified fixed-effects models assessing whether the effect of average occupational-level WFH on GHQ-12 scores varies by childcare status (living with children under 15 years old).

For men, no significant interaction between average occupational-level WFH and childcare status was found in either period.

For women, before March 2020, the interaction was significant, indicating that women with children under 15 experienced smaller mental health benefits from WFH compared to those without. However, the marginal effects of average occupational-level WFH presented in Figure 3 were not significant for women regardless of childcare responsibilities, nor did they differ significantly between groups, indicating no clear evidence of an effect within or between these subgroups.

[Table 2 about here]

[Figure 3 about here]

d. Variations in the effect of WFH on mental health with occupational class

Table 3 presents estimates from gender-stratified fixed-effects models assessing whether the effects of average occupational-level WFH on GHQ-12 scores vary by occupational class.

Before March 2020, among women, significant interaction terms indicate that the effects of average occupational-level WFH varied across occupational classes. Specifically, higher average occupational-level WFH led to poorer mental health for women in professional occupations, as indicated by negative main effects on the GHQ. The interaction effects for routine occupations are positive and significant, indicating that higher average occupational-level WFH benefited more women in routine jobs than those in professional roles. Figure 4 illustrates these patterns, showing significantly positive marginal effects of average occupational-level WFH on the GHQ for women in routine occupations, contrasted with negative marginal effects for those in professional roles. Analyses with the MCS-12 (Appendix F) again show similar patterns, confirming the heterogeneity of the effect of average occupational-level WFH on mental health across occupational classes.

From March 2020 onward, this trend reversed. Increases in average occupational-level WFH led to significant improvements in the mental health of women in professional occupations, as indicated by positive main effects on the GHQ. However, a significant interaction term shows that effects on the GHQ were significantly smaller for women in routine occupations. As shown in Figure 4, the marginal effect of occupational-level WFH on mental health was positive and statistically significant only for women in professional roles. Again, sensitivity analyses using the MCS-12 show similar results.

Among men, no significant interactions between average occupational-level WFH and occupational class were found in either period, suggesting that the mental health effects of WFH were homogenous across occupational classes.

[Table 3 about here]

[Figure 4 about here]

a. Discussion

In this paper, we provide a holistic picture of the impact of WFH on mental health in different subgroups of the UK population, before and after its widespread adoption during the COVID-19 pandemic. By focusing on within-individual changes and using an external occupational-level measure of WFH, we address selection issues and provide robust causal evidence on its mental health impact at the UK population level. While previous research has primarily focused on gender differences, we demonstrate that the mental health consequences of WFH vary at the intersection of gender, childcare responsibilities, and occupational class, and that these patterns have shifted as WFH became more institutionalised following the onset of the pandemic.

Our findings indeed reveal important heterogeneity in the mental health effects of WFH. Before March 2020, increases in the proportion of individuals WFH within men's occupations had significant positive effects on their mental health, and these effects were homogeneous across occupational classes. For women, variations in the level of WFH within their occupation did not have significant overall effects on mental health; however, analyses revealed heterogeneous effects across occupational classes. Specifically, increased WFH worsened mental health for women in higher-status positions—where WFH was more common—while it improved mental health for women in routine occupations—where WFH remained rare². The significant interaction between childcare status and WFH suggests that stressors at the work–family interface underlie these differences in the mental health effects of WFH.

² These findings contrast with some existing literature reporting positive effects of WFH on mental health for both women and men (Denzer and Grunau 2023), including studies using the same dataset but based on self-reported WFH measures (Li and Wang 2022). These differences may stem from our more rigorous controls for selection into WFH, but also from our occupational-level WFH measure, which captures more intensive and regular WFH use than previous studies. This suggests that high-intensity WFH may not provide mental health benefits for women, whereas more occasional WFH does, as proposed by Henke et al. (2016).

Several mechanisms may explain the patterns observed in how changes in WFH within occupations affected mental health across gender and occupational class before the pandemic.

For men, the positive effect of WFH likely reflects that they benefited from the increased job resources associated with WFH—such as job autonomy or schedule control—which helped them manage job demands while enabling greater and rewarding childcare involvement (Carlson et al. 2021). This aligns with evidence that, pre-pandemic, WFH was primarily accessible to workers with health limitations or in highly skilled positions, for whom it provided greater job autonomy (Milasi et al. 2021). It was also associated with a modest increase in unpaid labour, insufficient to generate role conflicts (Wheatley 2012).

For women in higher-status roles, the observed negative impact of WFH mental health echoes prior research showing that, unlike other flexible arrangements such as reduced hours or flexitime, WFH can lead to greater work–life conflicts (Russell et al. 2009, Yucel and Chung 2023). This suggests that for these higher-status women, who constituted most women WFH before the pandemic, WFH acted as ‘boundary-spanning demands’: it allowed them to maintain demanding, lucrative positions, often at the cost of increased work–family conflicts offsetting the benefits of job autonomy (Chung and van der Horst, 2018). These conflicts may have been further intensified by behaviours aimed at countering the flexibility stigma, such as extending work hours or multitasking (Stone and Hernandez 2013; Cech and Blair-Loy 2014).

By contrast, the positive impact of WFH we observe for women in lower-status occupations suggests that WFH may have functioned as a key ‘boundary-spanning resource’ for them. As these women have lower schedule control (Tausig et al. 2011; Damaske et al. 2016) and fewer resources to outsource unpaid care (Clawson and Gerstel 2014), WFH likely helped them better integrate work and family demands. This aligns with research integrating individuals’ occupational class into the job demands–resources framework, which finds that job resources

reduce work strain more for lower-status than higher-status workers (Koltai and Schieman 2015). Taken together, those results indicate that, at that time, WFH was not an effective workplace policy for alleviating psychological distress among higher-status women, who most frequently used it, though it did benefit those with fewer resources to manage competing work and family demands.

Interestingly, patterns of gender and occupational class differences in the mental health effects of WFH reversed from March 2020. Women, especially those in higher-status jobs, experienced significant improvements in their mental health as the level of WFH within their occupation increased, while no effect was observed for men. It is important to note that in the LFS, the increase in WFH from March 2020 reflects workers whose home-based work arrangements persisted beyond the initial months of the pandemic, rather than those who temporarily shifted to WFH. Then, these results may highlight the negative impact of returning to in-person work after the early pandemic period (Fan and Moen 2023a, Fan and Moen 2023b), rather than the benefits of WFH itself.

Still, there are several reasons why WFH may have become beneficial for the mental health of professional women in the post-pandemic period. First, as WFH became more normalized, the stigma surrounding flexible working, which disproportionately affects them, may have weakened (Blair-Loy 2003; Ridgeway and Corell 2024), along with the compensatory behaviours it triggers that tend to exacerbate job strain and work-family conflicts (Howcroft et al. 2024). Additionally, WFH-related strain may have been alleviated as organisational and policy support for remote work increased (Wheatley et al. 2021).

Conversely, several factors may explain why the mental health benefits of WFH observed pre-pandemic did not persist for men from March 2020 onwards. As noted above, WFH expanded to a broader range of occupations, and studies note that new forms of control emerged,

including bureaucratic procedures and technology-mediated managerial oversight (Fana et al., 2022). This likely reduced the autonomy WFH provided and, consequently, its associated mental health benefits. In addition, men WFH assumed greater unpaid labour during the pandemic, exposing some to heightened work–family conflicts (Collins et al. 2021). Similar dynamics may also explain why the pre-pandemic mental health benefits of WFH weakened for women in routine jobs. Once limited to very specific cases, WFH expanded to a wider range of jobs, while childcare demands increased and home-working conditions remained less supportive than in higher-status jobs (Ewers et al. 2023). Further research using post-pandemic data is needed to determine whether these trends have persisted and whether the normalisation and institutionalisation of WFH have durably reshaped its mental health effects.

Importantly, our results show that for both time periods, the mental health impact of WFH varies across individuals' social structural location, particularly at the intersection of gender and occupational class. They suggest that the stressors associated with WFH are not simply additive across axes of inequality; for example, WFH can serve as a compensatory resource for women in lower-status jobs that mitigates work-family conflicts and supports their mental health. This highlights the need for an intersectional approach to fully understand the mental health effects of WFH. Further research should explore other axes of inequality, such as race/ethnicity, age, and caregiving responsibilities beyond childcare, including eldercare. In additional analyses, we found no significant differences in the mental health effects of WFH by eldercare and race/ethnicity. However, this may reflect limitations in the measures used, indicating a need for more focused investigation in future studies.

This paper has several limitations. First, increases in WFH within occupations may coincide with other changes in job quality that also affect mental health, such as greater access to other flexible working arrangements (e.g., flexitime) or shifts in workload and managerial practices.

Therefore, our results may partly reflect the impact of these broader changes. However, supplementary analyses showed no significant association between our occupational-level WFH measure and self-reported use of other flexible working arrangements, suggesting these factors are unlikely to confound our findings.

Second, as noted above, shifts in WFH effects before and after the pandemic likely reflect changes in both its users and its implementation—factors our design cannot fully capture. Quasi-experimental studies, for instance, exploiting policy reforms expanding WFH access, are needed to better isolate its impact from concurrent shifts in workforce and job characteristics (Xue et al. 2025). Nonetheless, as the effects of WFH seem context-dependent, such causal studies should be complemented by broader analyses like the present one.

Third, while we hypothesise that WFH affects mental health through job strain and work-family conflict, our analyses do not directly test these pathways using mediation analysis or specific measures of these stressors. Nor do we examine potential moderators, such as the frequency of WFH or the quality of the WFH environment. This is due to the scope of our paper, which prioritises addressing reverse causality by relying on an external measure of WFH, since individual WFH use, but also self-reported stressors, are influenced by mental health. Although this approach strengthens the causal interpretation of our findings, it restricts our ability to explore mediating mechanisms. Combining it with descriptive approaches is necessary to fully understand how WFH intersects with mental health inequalities.

Despite these limitations, this study provides robust evidence of the potential mental health benefits of WFH, whilst also showing the conditions under which these benefits may be lost. For both men and women, the mental health outcomes of WFH depend on circumstances tied to their social positions as well as on how organisations and policies manage the demands of flexible work. As such, WFH can be a key resource to mitigate the growing mental health

challenges faced by today's workforce. However, realising this potential requires WFH to be implemented alongside supportive organisational practices and broader social policies that help workers manage the intensifying demands of both paid work and family life.

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TABLES

Table 1: Effects of average occupational-level WFH on GHQ-12, with gender interaction (two-way fixed effects models)

Period	Term	Estimate	95% Confidence Interval	P-Value
January 2009 – February 2020	Mean WFH	0.682	[0.183; 1.182]	0.008
	Mean WFH x Women	-0.861	[-1.627; -0.094]	0.028

	<i>Number of respondent-wave observations: 169,309 · Number of unique respondents: 39,863 · Adjusted R-squared: 0.423 · Within R-squared: 0.004</i>			
March 2020 – October 2023	Mean WFH	-0.316	[-1.106; 0.475]	0.433
	Mean WFH x Women	1.415	[0.267; 2.563]	0.016
	<i>Number of respondent-wave observations: 37,159 · Number of unique respondents: 18,441 · Adjusted R-squared: 0.510 · Within R-squared: 0.004</i>			

Table 2: Effects of average occupational-level WFH on GHQ-12, with childcare status interaction (two-way fixed effects models)

Period	Term	Estimate	95% Confidence Interval	P-Value
Women				
January 2009 - February 2020	Children < 15 yo (vs. no children)	0.076	[-0.084; 0.235]	0.353
	Mean WFH	0.309	[-0.391; 1.01]	0.386

	Mean WFH x Children < 15 yo (vs. no children)	-0.962	[-1.905; -0.019]	0.046
	Number of respondent-wave observations: 92,743 · Number of respondents: 21,435 · Adjusted R ² : 0.414 · Within R ² : 0.002			
March 2020 - October 2023	Children < 15 yo (vs. no children)	-0.254	[-1.011; 0.502]	0.509
	Mean WFH	0.848	[-0.461; 2.157]	0.203
	Mean WFH x Children < 15 yo (vs. no children)	0.037	[-1.68; 1.754]	0.966
	Number of respondent-wave observations: 19,897 · Number of respondents: 9,896 · Adjusted R ² : 0.482 · Within R ² : 0.004			
Men				
January 2009 - February 2020	Children < 15 yo (vs. no children)	-0.156	[-0.351; 0.038]	0.115
	Mean WFH	0.728	[0.091; 1.366]	0.025
	Mean WFH x Children < 15 yo (vs. no children)	-0.137	[-1.023; 0.748]	0.76
	Number of respondent-wave observations: 79,577 · Number of respondents: 18,967 · Adjusted R ² : 0.419 · Within R ² : 0.005			

March 2020 - October 2023	Children < 15 yo (vs. no children)	-0.068	[-0.773; 0.637]	0.849
	Mean WFH	0.07	[-0.863; 1.004]	0.882
	Mean WFH x Children < 15 yo (vs. no children)	-0.399	[-1.564; 0.767]	0.501
	Number of respondent-wave observations: 17,262 · Number of respondents: 8,546 · Adjusted R ² : 0.532 · Within R ² : 0.003			

Table 3: Effects of average occupational-level WFH on GHQ-12, with occupational class interaction (two-way fixed effects models)

Period	Term	Estimate	95% Confidence Interval	P-Value
Women				
January 2009 - February	Mean WFH	-0.735	[-1.459; -0.011]	0.047

	Mean WFH x Intermediate (vs. Professional)	0.627	[-0.482; 1.736]	0.267
	Mean WFH x Routine (vs. Professional)	2.507	[1.398; 3.617]	0
	<i>Number of respondent-wave observations: 91,817 · Number of respondents: 21,415 · Adjusted R²: 0.414 · Within R²: 0.003</i>			
March 2020 - October 2023	Mean WFH	1.071	[0.016; 2.127]	0.047
	Mean WFH x Intermediate (vs. Professional)	-0.577	[-1.949; 0.795]	0.409
	Mean WFH x Routine (vs. Professional)	-2.955	[-5.927; 0.016]	0.051
	<i>Number of respondent-wave observations: 19,448 · Number of respondents: 9,795 · Adjusted R²: 0.484 · Within R²: 0.005</i>			
Men				
January 2009 - February 2020	Mean WFH	0.485	[-0.200; 1.17]	0.165
	Mean WFH x Intermediate (vs. Professional)	-0.047	[-1.108; 1.013]	0.93

	Mean WFH x Routine (vs. Professional)	0.497	[-0.373; 1.366]	0.262
	<i>Number of respondent-wave observations: 78,762 · Number of respondents: 18,946 · Adjusted R²: 0.421 · Within R²: 0.005</i>			
March 2020 - October 2023	Mean WFH	0.019	[-0.879; 0.916]	0.968
	Mean WFH x Intermediate (vs. Professional)	-0.401	[-1.961; 1.16]	0.614
	Mean WFH x Routine (vs. Professional)	2.111	[-0.386; 4.608]	0.097
	<i>Number of respondent-wave observations: 16,810 · Number of respondents: 8,435 · Adjusted R²: 0.536 · Within R²: 0.006</i>			

FIGURES

Figure 1: Proportion of individuals reporting access to and use of WFH in their main occupation in UKHLS, and proportion of individuals reporting being mainly WFH in UK LFS, by year

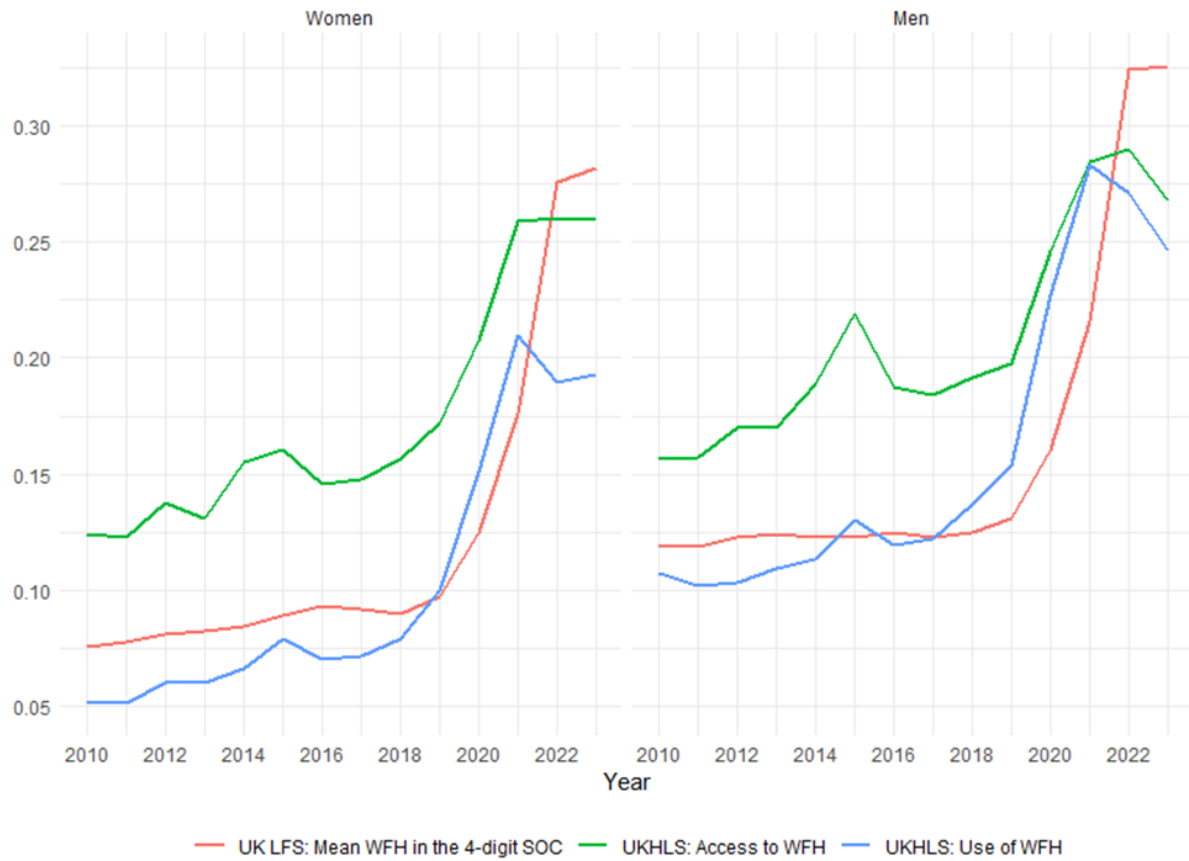
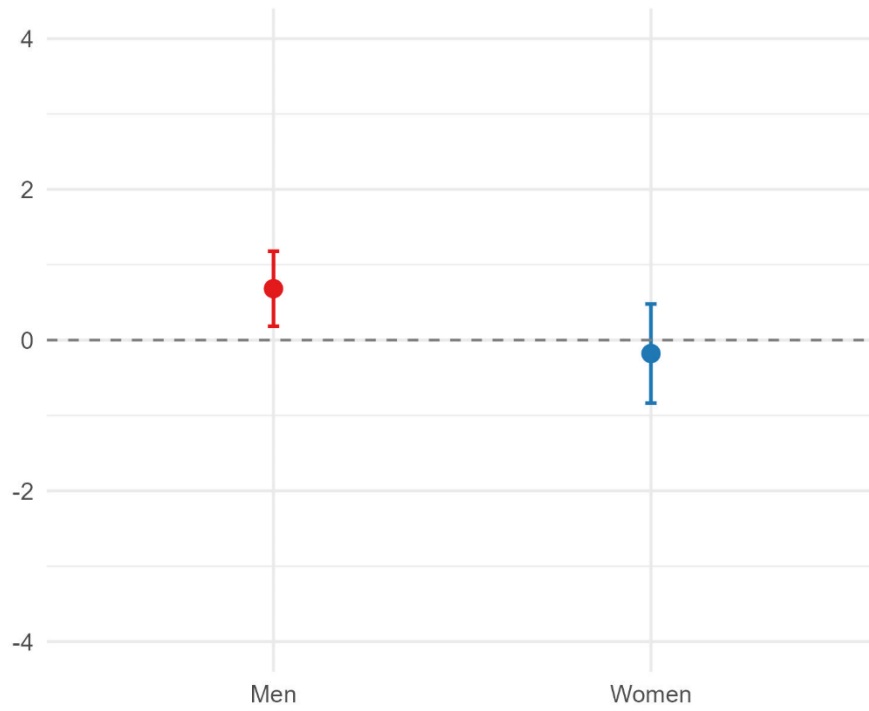


Figure 2: Marginal effects of average occupational-level WFH on GHQ-12, by gender, before and from March 2020

January 2009 – February 2020



March 2020 – October 2023

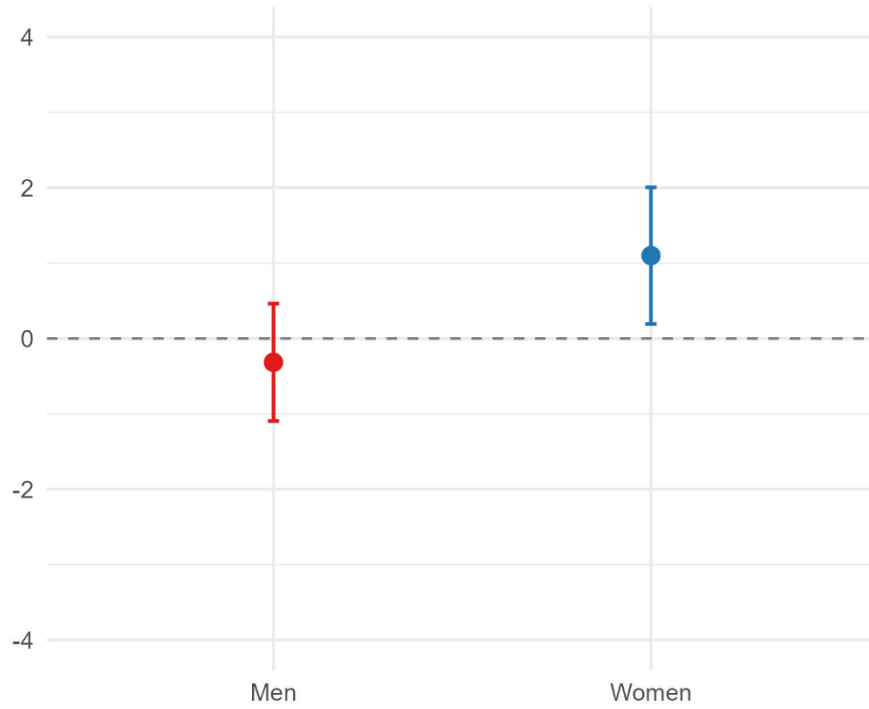


Figure 3: Marginal effects of average occupational-level WFH on GHQ-12, by gender and childcare status, before and from March 2020

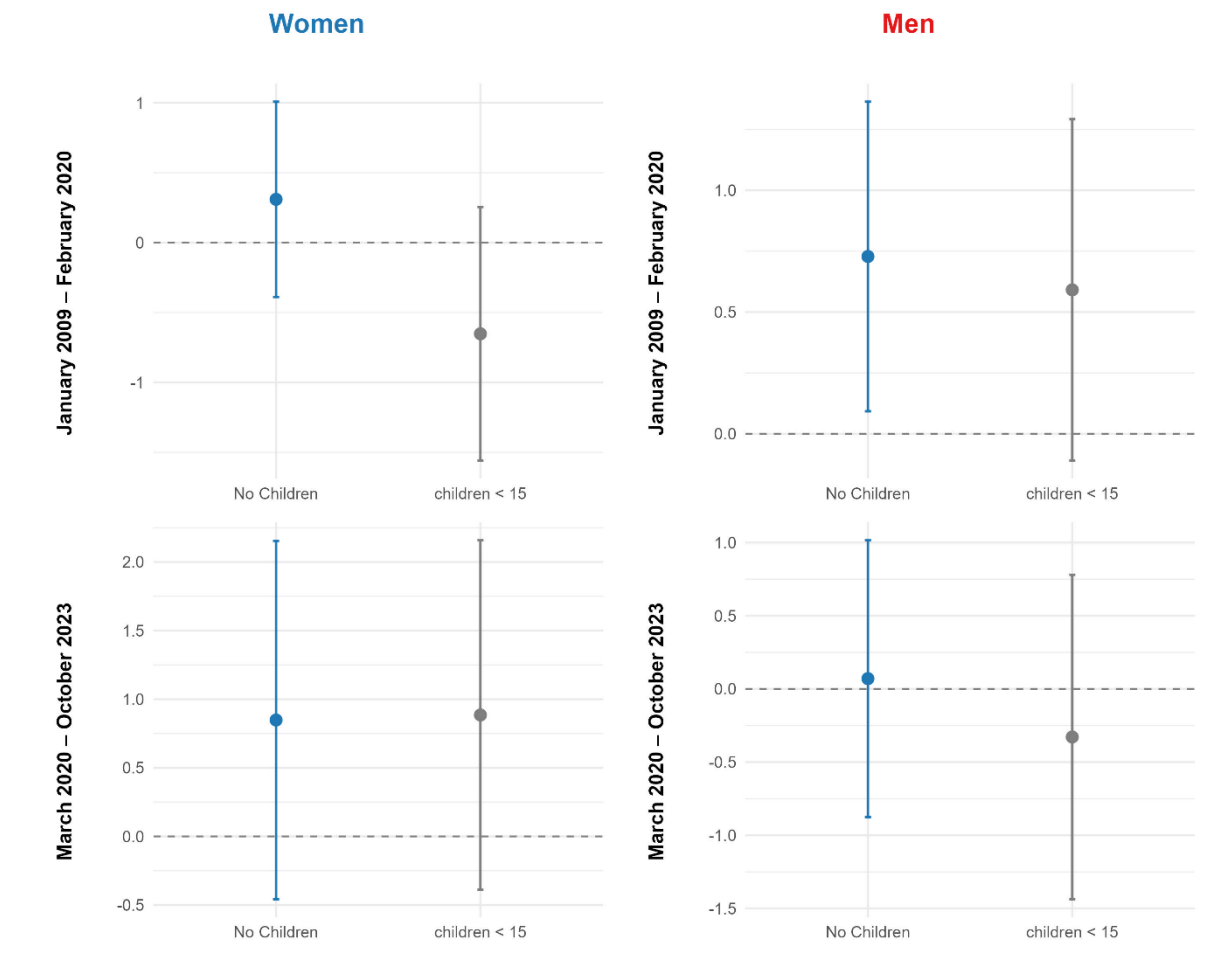
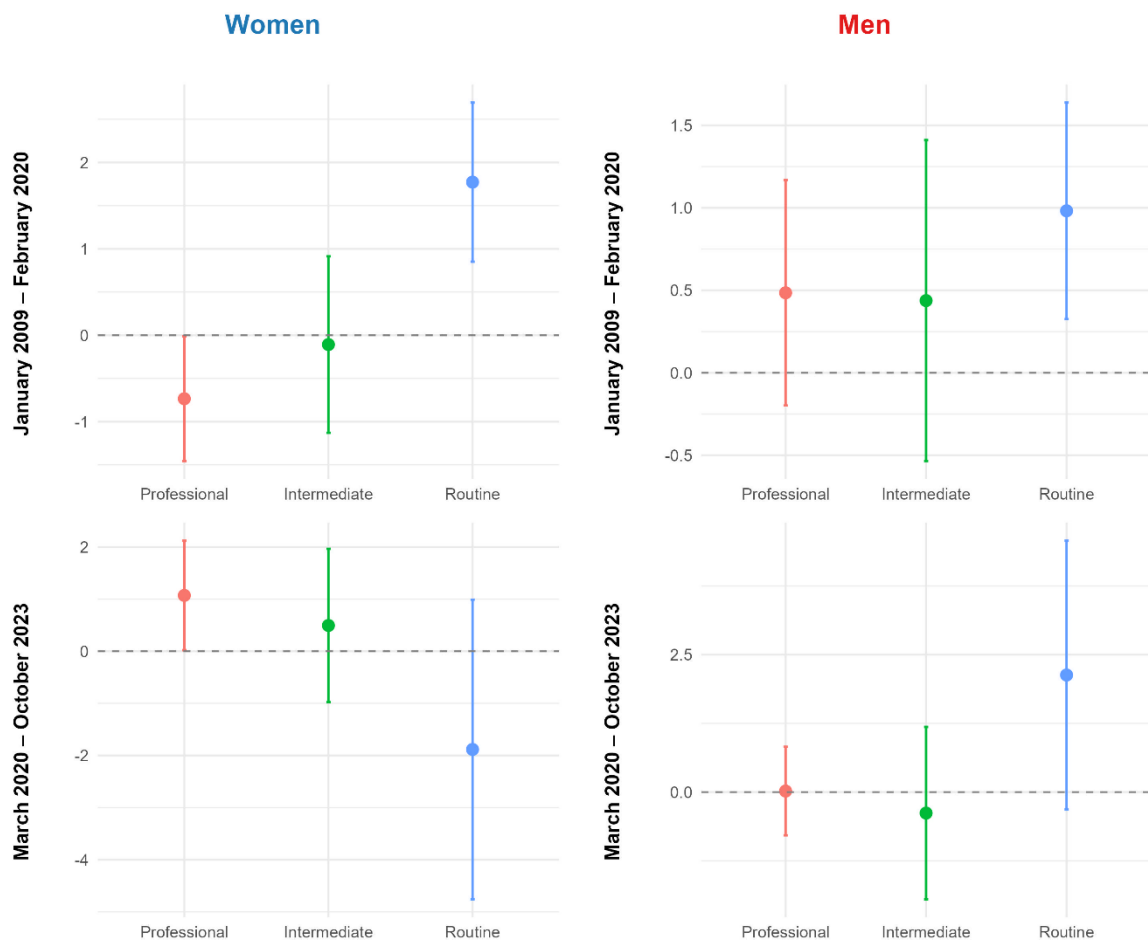


Figure 4: Marginal effects of average occupational-level WFH on GHQ-12, by gender and occupational class, before and from March 2020



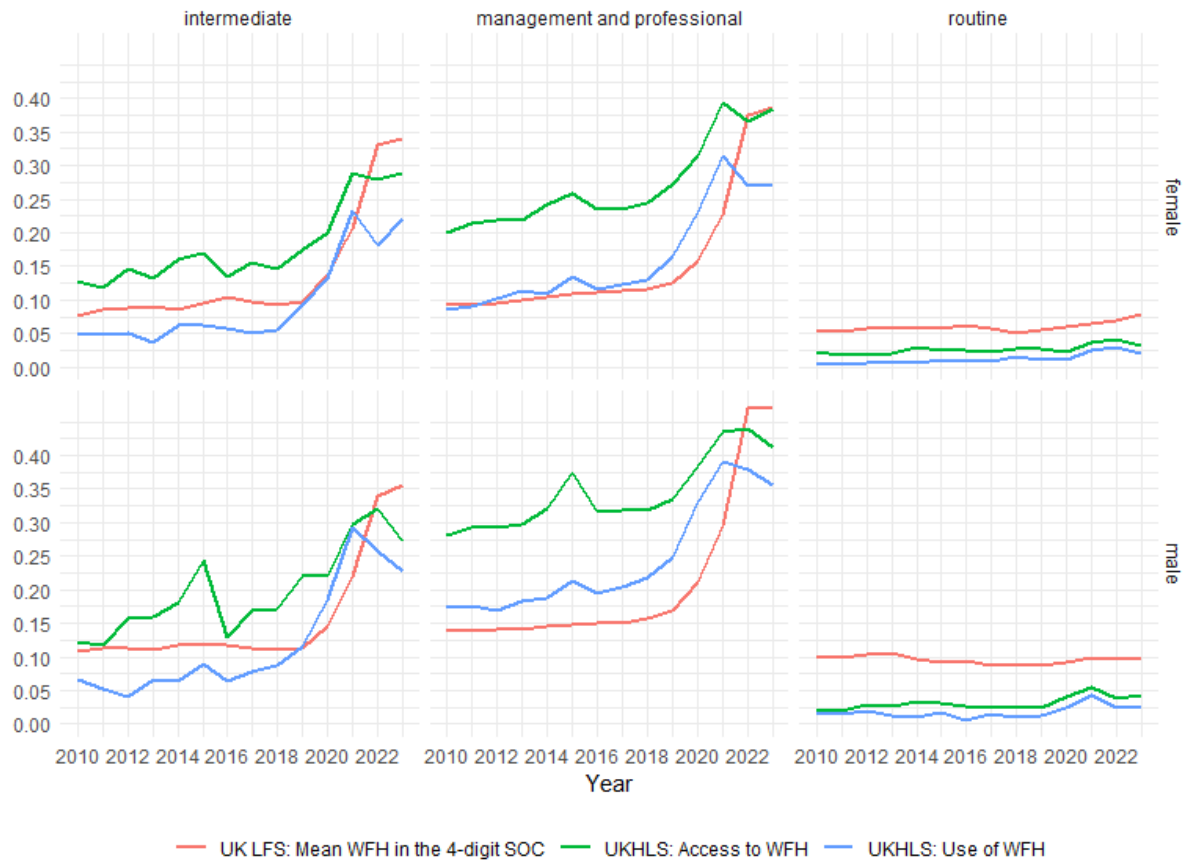
Appendices

Appendix A: Sociodemographic and health characteristics of the sample

Characteristics	Women (N = 115,604)	Men (N = 99,572)	p-value
Age			<0.001
Mean (SD)	41.22 (11.22)	42.14 (11.93)	
Marital status			<0.001
Married	60,615 (52.43%)	58,858 (59.11%)	
Separated / Divorced / Widowed	16,615 (14.37%)	8,449 (8.49%)	
Single	38,374 (33.19%)	32,265 (32.40%)	
Monthly household income (net, real)			<0.001
Mean (SD)	4,678.84 (5,781.71)	4,774.39 (4,490.50)	
Occupational class (NS-SEC, 3-category)			<0.001
Intermediate occupations	23,355 (20.79%)	11,306 (11.71%)	
Managerial and professional occupations	51,819 (46.12%)	47,589 (49.28%)	
Routine and manual occupations	37,182 (33.09%)	37,667 (39.01%)	
Weekly working hours			<0.001
Mean (SD)	30.31 (10.17)	37.67 (8.57)	
Missing	2,903	2,664	

Average occupational-level WFH			<0.001
Mean (SD)	0.11 (0.13)	0.15 (0.15)	
GHQ-12			<0.001
Mean (SD)	11.48 (5.40)	10.36 (4.77)	
Child under 15 years in the household			<0.001
Yes	48,093 (41.62%)	40,310 (40.51%)	
No	67,448 (58.38%)	59,191 (59.49%)	
Providing adult care			<0.001
Yes	20,077 (17.37%)	11,966 (12.02%)	
No	95,527 (82.63%)	87,606 (87.98%)	
Number of person-wave observations	115,604	99,572	
Number of unique respondents	24,160	21,404	
¹ n (%)			
² Wilcoxon rank sum test; Pearson's Chi-squared test			

Appendix B: Mean level of individuals reporting access and use to WFH in their main occupation in UKHLS, and mean level of individuals reporting being mainly WFH in UK LFS, by year



Appendix C: Associations between average occupational-level WFH and self-reported access and use of WFH (Two-way fixed effects models)

Outcome	Predictor	Estimate	95% Confidence Interval	p-value
Access to WFH (all waves)	Mean WFH	2.024	[1.514, 2.533]	<0.001
	Mean WFH × Men	−1.048	[−1.615, −0.481]	<0.001
Access to WFH (pre-March 2020)	Mean WFH	1.998	[0.929, 3.066]	<0.001
	Mean WFH × Men	−1.502	[−2.960, −0.044]	0.044
Use of WFH (all waves)	Mean WFH	1.642	[0.956, 2.328]	<0.001
	Mean WFH × Men	−0.383	[−1.108, 0.342]	0.300
Use of WFH (pre-March 2020)	Mean WFH	1.925	[0.409, 3.440]	0.013
	Mean WFH × Men	−1.230	[−3.387, 0.928]	0.264

Appendix D: Effects of average occupational-level WFH on SF-12, with gender interaction (two-way fixed effects models)

Period	Term	Estimate	95% Confidence Interval	P-Value
January 2009 – February 2020	Mean WFH	0.773	[-0.093; 1.640]	0.080
	Mean WFH x Women	-0.988	[-2.359; 0.383]	0.157
	<i>Number of respondent-wave observations: 166,429 · Number of unique respondents: 39,279 · Adjusted R-squared: 0.471 · Within R-squared: 0.005</i>			
March 2020 – October 2023	Mean WFH	-1.048	[-2.425; 0.33]	0.136
	Mean WFH x Women	2.834	[1.057; 4.61]	0.002
	<i>Number of respondent-wave observations: 36,896 · Number of unique respondents: 18,363 · Adjusted R-squared: 0.587 · Within R-squared: 0.004</i>			

Appendix E: Effects of average occupational-level WFH on SF-12, with childcare status interaction (two-way fixed effects models)

Period	Term	Estimate	95% Confidence Interval	P-Value
Women				
January 2009 - February 2020	Children < 15 yo (vs. no children)	0.075	[-0.194; 0.345]	0.582
	Mean WFH	0.555	[-0.83; 1.939]	0.431
	Mean WFH x Children < 15 yo (vs. no children)	-1.346	[-3.083; 0.39]	0.128
	Number of respondent-wave observations: 91,171 · Number of respondents: 21,153 · Adjusted R ² : 0.458 · Within R ² : 0.004			
March 2020 - October 2023	Children < 15 yo (vs. no children)	-0.233	[-1.299; 0.833]	0.668
	Mean WFH	1.638	[-0.386; 3.662]	0.112
	Mean WFH x Children < 15 yo (vs. no children)	0.234	[-2.657; 3.126]	0.873
	Number of respondent-wave observations: 19,750 · Number of respondents: 9,851 · Adjusted R ² : 0.564 · Within R ² : 0.002			
Men				
January 2009 - February 2020	Children < 15 yo (vs. no children)	-0.296	[-0.644; 0.051]	0.094
	Mean WFH	1.464	[0.354; 2.573]	0.01
	Mean WFH x Children < 15	-1.288	[-2.932; 0.355]	0.124

	yo (vs. no children)			
	<i>Number of respondent-wave observations: 78,130 · Number of respondents: 18,644 · Adjusted R²: 0.473 · Within R²: 0.006</i>			
March 2020 - October 2023	Children < 15 yo (vs. no children)	-0.594	[-1.871; 0.682]	0.361
	Mean WFH	-1.172	[-2.903; 0.559]	0.184
	Mean WFH x Children < 15 yo (vs. no children)	0.405	[-1.643; 2.452]	0.698
	<i>Number of respondent-wave observations: 17,146 · Number of respondents: 8,513 · Adjusted R²: 0.598 · Within R²: 0.005</i>			

Appendix F: Effects of average occupational-level WFH on GHQ-12, with occupational class interaction (two-way fixed effects models)

Period	Term	Estimate	95% Confidence Interval	P-Value
Women				
January 2009 - February 2020	Mean WFH	-1.013	[-2.221; 0.194]	0.1
	Mean WFH x Intermediate (vs. Professional)	0.084	[-1.859; 2.028]	0.932
	Mean WFH x Routine (vs. Professional)	4.59	[2.407; 6.773]	0
	<i>Number of respondent-wave observations: 90,256 · Number of respondents: 21,130 · Adjusted R²: 0.459 · Within R²: 0.004</i>			
March 2020 - October 2023	Mean WFH	2.022	[0.396; 3.648]	0.015
	Mean WFH x Intermediate (vs. Professional)	-1.417	[-3.784; 0.95]	0.24
	Mean WFH x Routine (vs. Professional)	-1.84	[-7.008; 3.327]	0.484
	<i>Number of respondent-wave observations: 19,310 · Number of respondents: 9,753 · Adjusted R²: 0.567 · Within R²: 0.004</i>			
Men				
January 2009 - February 2020	Mean WFH	0.524	[-0.604; 1.652]	0.361

	Mean WFH x Intermediate (vs. Professional)	1.146	[-0.448; 2.739]	0.158
	Mean WFH x Routine (vs. Professional)	0.663	[-0.851; 2.177]	0.39
	<i>Number of respondent-wave observations: 77,328 · Number of respondents: 18,622 · Adjusted R²: 0.473 · Within R²: 0.006</i>			
March 2020 - October 2023	Mean WFH	-0.985	[-2.544; 0.574]	0.215
	Mean WFH x Intermediate (vs. Professional)	0.196	[-2.245; 2.637]	0.875
	Mean WFH x Routine (vs. Professional)	1.501	[-3.084; 6.086]	0.52
	<i>Number of respondent-wave observations: 16,698 · Number of respondents: 8,404 · Adjusted R²: 0.604 · Within R²: 0.006</i>			