Who Wants to Work from Home?

Stated Preferences and Realized Access in the German Labour Market, 2009–2014

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Abstract

Most research on remote work asks who can work from home, but rarely who actually wants to. Without a direct measure of preferences, low WFH rates and unmet demand are indistinguishable. This paper uses German Socio-Economic Panel (SOEP) data for 2009–2014 to study both questions in a two-stage framework. The first stage estimates which characteristics predict WFH demand among full-time employees in teleworkable occupations. The second stage estimates which characteristics predict whether that demand is actually realized, restricting the sample to workers who stated a positive preference. Both stages use pooled logit and probit models alongside a Mundlak correlated random effects probit.

Women are about 3 percentage points more likely to want WFH but 5.6 to 6.7 percentage points less likely to obtain it, a gap that holds after within-individual estimation. Workers with a direct migration background are about 14 percentage points less likely to want WFH, yet face no access disadvantage conditional on having stated a preference. Education and income are positively associated with WFH outcomes in the pooled models, but both largely reflect occupational sorting once individual heterogeneity is controlled for. The results show that WFH inequality takes different forms across groups, with implications for how access-side policies such as a statutory right to WFH should be targeted.

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1. Introduction

Remote work has grown from a niche arrangement into a recognised feature of the labour market. The share of employees working at least occasionally from home has expanded substantially over recent decades, reaching approximately 24% by 2023 (Destasis 2023). What was once confined to specific occupational groups has become a common element of the employment relationship, raising questions about who gains access to this arrangement and on what terms.

The existing literature has established that access to remote work is unequally distributed along socioeconomic lines. Yasenov (2020) finds that lower-wage workers are up to three times less likely to be in teleworkable occupations, with lower-educated workers, younger adults, ethnic minorities, and immigrants disproportionately concentrated in non-remote jobs. Bonacini et al. (2021) project that an expansion of WFH (work from home) feasibility would disproportionately benefit male, older, highly educated, and high-paid employees, risking an exacerbation of pre-existing labour market inequalities. Taken together, this literature treats the inability to work from home as the primary source of disadvantage and interprets any demographic gap in remote work rates as evidence of unequal access.

What this literature rarely asks is who actually wants to work from home. Without a direct measure of preferences, low WFH rates and unmet demand are indistinguishable: a worker who does not work from home may face a genuine access barrier, or may simply not want the arrangement. There are three key reasons why this distinction matters.

First, it is an equity issue. The welfare-relevant inequality is not who works from home but who wants to and cannot. A large preference-realization gap for a particular socioeconomic or demographic group constitutes evidence of constrained exclusion, not voluntary non-participation, and represents an inequality that raw uptake statistics cannot reveal.

Second, it bears directly on policy design. Proposals for a statutory right to work from home have surfaced repeatedly in European legislative debates (SPD 2020; Milders 2022; Citizens Information 2024; UNECE 2020), reflecting sustained pressure from workers and unions in remote-compatible sectors. Such a right is only a meaningful intervention where the gap between stated preferences and realized access is large; where WFH rates are low because workers do not want the arrangement, a statutory entitlement provides no welfare gain. Conversely, if stated preferences are themselves concentrated among groups already advantaged in terms of education, income, or domestic circumstances, a universal right could produce regressive distributional outcomes, even if enacted with universalist intent.

Third, the gap is interpretable as a labour market imperfection. Under the theory of compensating differentials, workers in a competitive market trade wages for job amenities, and employers offering worse amenities must compensate with higher pay. Remote work is one such amenity, and workers willing to accept lower wages in exchange for it should, in equilibrium, be able to obtain it. A systematic failure to realize stated preferences therefore points to employer-side bargaining power over the terms of the employment relationship, not merely to individual misfortune.

By directly measuring stated preferences for WFH alongside realized outcomes, this paper addresses two related questions. First, which socioeconomic and demographic characteristics predict whether a worker prefers to work from home? Second, among workers who express the preference to work from home, which characteristics predict whether it is actually realized? The gap between these two questions, who wants WFH and who gets it, is the central quantity of interest.

The main findings of this paper are as follows. Among workers in teleworkable occupations, women are about 3 percentage points more likely to state a positive WFH preference, while workers with a direct migration background are about 14 percentage points less likely to do so. Among workers who stated a preference, women are 5.6 to 6.7 percentage points less likely to actually work from home, an estimate that holds after controlling for individual fixed heterogeneity. Workers with a migration background, by contrast, face no access disadvantage once they have stated a preference. Education and income are positively associated with WFH outcomes in the pooled models. For income, the correlated random effects specifications suggest this largely reflects sorting into WFH-compatible roles rather than within-job barriers. For education, the panel cannot separately identify a within-person effect, as within-individual variation in educational attainment is near-zero over the observation window. Taken together, the results show that WFH inequality takes different forms across groups: a genuine access barrier for women, a demand-side gap for migrants, and a sorting effect for income.

2. Literature Review

2.1 Who Can Work from Home?

A foundational strand of the WFH literature asks a prior question before anything about preferences or outcomes: which jobs are structurally compatible with remote work? Dingel and Neiman (2020) provide the most widely cited answer, classifying occupations using O*NET task descriptors and finding that approximately 37% of jobs in the United States can be performed entirely from home, jobs that account for 46% of total US wages. Applying the same classification to 85 other countries, they show that the share of teleworkable jobs falls sharply with national income, embedding the access question in a broader inequality framework from the outset.

Subsequent work has broadly replicated this figure while refining the methodology. Holgersen et al. (2021) use ISCO-08 task descriptions evaluated by Amazon Mechanical Turk respondents to estimate that roughly 38–39% of Norwegian jobs are teleworkable, and stress that the potential for remote work is unevenly distributed: workers already at a disadvantage in the labour market are systematically less likely to hold occupations that can be performed from home. Yasenov (2020) reaches a similar conclusion for the United States, finding that lower-wage workers are up to three times less likely to be able to work remotely than higher-wage workers, with lower-educated workers, younger adults, ethnic minorities, and immigrants disproportionately concentrated in non-teleworkable occupations.

The 37% estimate for the United States has been replicated across European and global contexts. Sostero et al. (2020) construct teleworkability indices for 121 ISCO occupation groups using European task data and find that approximately 37% of EU dependent employment is technically teleworkable, but with a steep socioeconomic gradient: while 74% of jobs in the top wage quintile are teleworkable, only 3% of those in the bottom quintile are. Sostero et al. (2023) further show that pre-pandemic uptake among white-collar workers was well below this technical potential, attributing the gap to hierarchical norms and management culture rather than task constraints, a finding that foreshadows the role of employer behaviour explored later in this thesis. Bamieh and Ziegler (2022) provide direct evidence of the employer-side shift during COVID-19: using vacancy postings from the Austrian Public Employment Service, they find that employers became two to three times more likely to explicitly offer teleworking in job advertisements during the first year of the pandemic, with the change concentrated in higher-skilled roles requiring secondary or university education.

Beyond high-income economies, the teleworkability gap widens sharply. Sanchez et al. (2020) apply a modified version of the Dingel and Neiman (2020) framework to 107 countries and find that, once residential internet access is accounted for, only 18.7% of global jobs are feasibly teleworkable. High-income countries can support approximately one in three jobs remotely, while low-income countries are limited to roughly one in twenty-six; failing to account for internet access overestimates developing-economy resilience by a factor of four. Berg et al. (2020) reach a similar conclusion using the Delphi method across 89 countries, estimating 27% teleworkability in high-income countries against 12% in low-income ones. Gottlieb et al. (2021) corroborate this for urban workers in ten low- and middle-income countries, finding fewer than 10% of jobs teleworkable due to task content and limited information and communication technology (ICT) usage.

The COVID-19 pandemic provided a natural test of these predictions at scale. Montenovo et al. (2020) link CPS employment data from February to May 2020 with O*NET occupation scores and find that remote work capacity was a critical protective factor: workers in teleworkable occupations were significantly less likely to experience unemployment during the early pandemic. Importantly, a large share of demographic gaps in job losses (by race, ethnicity, and education) can be explained by pre-pandemic sorting into occupations with different remote work potential, rather than by direct discrimination. Beland et al. (2023) corroborate this using a longer window, showing that workers in occupations requiring physical proximity to others were hit hardest, while those able to work remotely and those classified as essential were comparatively shielded.

Taken together, this literature establishes that access to remote work is deeply unequal and largely determined by occupational class before any individual characteristics come into play. However, all of these studies classify teleworkability at the occupation level and largely leave open whether individual workers within teleworkable occupations actually have access, nor whether they want it.

2.2 Social Inequality and Remote Work Access

Beyond occupational classification, a related body of work examines how access to WFH maps onto existing socioeconomic fault lines within countries. Bonacini et al. (2021) use influence function regression on Italian employees to project how a sustained increase in WFH feasibility would affect the labour income distribution. Their central finding is that the gains would be unequally distributed: male, older, highly educated, and high-paid employees would benefit most, while the expansion of remote work risks exacerbating pre-existing labour market inequalities unless actively regulated.

The gender dimension is explored explicitly by Alon et al. (2020), who argue that the pandemic recession differed from prior downturns precisely because it hit female-dominated sectors hard while simultaneously collapsing childcare infrastructure. The resulting pressure fell disproportionately on working mothers, with potentially persistent consequences

Table 1: Teleworkability and realised WFH prevalence across selected studies.

Study	Region / Group	Period	Share (%)
Panel A: Aggregate teleworkability			
Dingel and Neiman (2020)	United States	2019	37
Holgersen et al. (2021)	Norway	2019	38–39
Sostero et al. (2020)	European Union	2020	37
Sanchez et al. (2020)	Global†	2019	18.7
Berg et al. (2020)	High-income	2020	27
Berg et al. (2020)	Middle-income	2020	16
Berg et al. (2020)	Low-income	2020	12
Gottlieb et al. (2021)	Developing	2012–13	<10
Panel B: Distributional inequality in teleworkability			
Dingel and Neiman (2020)	United States (wage share)	2019	46
Sostero et al. (2020)	European Union (top quintile)	2020	74
Sostero et al. (2020)	European Union (bottom quintile)	2020	3
Panel C: Realised WFH prevalence			
Barrero et al. (2023)	United States	2019	7
Barrero et al. (2023)	United States	2023	28

Notes: Panels A and B report the share of jobs technically performable from home based on task classification. †Adjusted for residential internet access. The 46% figure for Dingel and Neiman (2020) in Panel B denotes the share of total wages earned in WFH-compatible occupations; the remaining Panel B figures report the share of jobs classified as teleworkable within each wage quintile. Panel C reports the share of full working days actually spent at home.

due to high returns to labour market experience. These findings are directly relevant to the preference stage of this thesis: to the extent that WFH increases the domestic burden on women, stated preferences for remote work among mothers may reflect constrained choices rather than genuine demand.

2.3 Work from Home and Subjective Well-Being

A third strand of the literature moves from access to outcomes, asking how working from home actually affects workers' well-being. The evidence is surprisingly mixed and strongly heterogeneous by gender and family structure, a pattern that bears directly on who might want WFH when it is voluntary.

Senik et al. (2024) provide the most directly relevant study for the present thesis. Using the German Socio-Economic Panel (SOEP) and a difference-in-differences design, they find that teleworking had a negative average effect on life satisfaction over the first two years of the pandemic. However, this average conceals substantial heterogeneity: the negative effect is concentrated among unmarried men and women with school-age children, while women with children below school age report higher life satisfaction under

WFH. Notably, the negative effect for mothers of school-age children disappears by 2021, suggesting adaptation. As the only SOEP-based study of WFH well-being covering the mandate period, their results provide important context for interpreting the preference data used in this thesis.

Drawing on German data, Bellmann and Hübler (2020) use three waves of the Linked Personnel Panel from 2012 to 2016 and find that the introduction of remote work provides a temporary boost to job satisfaction that does not persist over time. Remote work also does not automatically improve work-life balance, and can disrupt it when driven by job demands rather than worker preference. Formally contracted arrangements yield higher satisfaction than informal ones, and the termination of existing WFH access significantly worsens work-life balance. The manner in which remote work is introduced and subsequently regulated therefore matters as much as its mere availability.

Gueguen and Senik (2023) reach a similar conclusion using UK Household Longitudinal Survey data and an individual fixed-effects design that exploits multiple entries into WFH. Full-time remote work worsens mental health on average, but this effect follows a dynamic pattern: the initial deterioration gives way to adaptation after a couple of months. The negative impact is most pronounced for women with children, particularly in the early months, which the authors associate with the home-schooling burden. Schifano (2023) draw on the COME-HERE panel across five European countries and find that WFH is associated with lower well-being than on-site work, with the penalty larger for older workers, the highly educated, those with young children, and those in overcrowded housing.

Not all evidence points in the same direction. Perelman et al. (2021) use SHARE Wave 8 data on Europeans aged 50–65 and find that, once the stringency of national lockdown policies is controlled for, WFH is not a significant driver of mental health deterioration; the decline in well-being is largely attributable to the pandemic context rather than the work arrangement itself. Xiao et al. (2021) survey US office workers and identify a more granular picture: decreased well-being is associated with reduced physical activity, poor nutrition, limited colleague communication, and the presence of young children, while a dedicated workspace and satisfying indoor conditions are protective.

The gendered dynamics of WFH are further documented by Lyttelton et al. (2023) using the American Time Use Survey. They find that remote work during the pandemic did not increase primary childcare time but led to a large rise in simultaneous supervision while working, a form of multitasking that fell disproportionately on mothers, who also experienced more work schedule disruptions than fathers. Cavapozzi (2022) provides causal estimates from SHARE data, finding that remote work increases the probability of reporting depression, with larger effects for women, workers with children at home, and single workers.

The mechanisms linking WFH conditions to job satisfaction are examined by García-Salirrosas et al. (2023) and Ploszaj et al. (2025) using survey data from remote workers in Latin America. Both studies identify work-life balance as a critical mediating pathway: supportive supervisory behaviours and favourable work-from-home conditions improve satisfaction primarily by enhancing work-life balance (WLB) and facilitating positive work-to-family spillover. Ploszaj et al. (2025) additionally document a gender asymmetry in this mediation: female satisfaction depends predominantly on stress reduction, whereas male satisfaction is achieved through a combination of reduced stress and improved WLB. While the cultural context differs from Germany, these findings reinforce the broader point that WFH's effect on satisfaction is not direct but mediated by domestic and organisational conditions. This mediation bears on whose preferences for WFH are most sensitive to the home environment they would actually face.

One finding in this literature is directly relevant to the preference stage of this thesis: Felstead and Reuschke (2020) report that, despite an initial decline in mental health during the UK's first lockdown, nearly 90% of workers who had shifted to WFH expressed a desire to continue doing so in the future, even as subjective well-being was, on average, lower. This divergence between experienced well-being and stated preferences motivates the approach taken here: rather than inferring preferences for WFH from outcomes, this thesis measures it directly using the self-reported preference variable available in the SOEP.

2.4 Productivity, Wages, and Work Burden of WFH

The impact of WFH on productivity depends critically on whether one estimates the average treatment effect of the environment or accounts for the selection of workers into remote arrangements. Bloom et al. (2015) provide the most widely cited causal evidence: a randomised controlled trial at a large Chinese travel firm found that home workers raised performance by 13%, with gains attributed to fewer breaks, fewer sick days, and a quieter environment. Yet home workers were promoted at half the rate of equally productive office colleagues, indicating that remote work can impose career penalties even when it raises output. When employees were subsequently allowed to self-select their work location, productivity gains nearly doubled to 22%, as workers who found themselves less productive at home voluntarily returned to the office, driven by preference-based sorting rather than the arrangement itself. Choudhury et al. (2021) extend this logic, finding that a transition from standard WFH to fully location-flexible work-from-anywhere at the United States Patent and Trademark Office raised patent examiner output by a further 4.4% with no quality deterioration.

More recent evidence suggests that selection into remote work can be at least as consequential as the environment. Emanuel and Harrington (2024) exploit pandemic-driven office closures at a Fortune 500 US call center and decompose a 12-percentage-point pre-pandemic productivity gap between remote and on-site workers: approximately 4 percentage points are attributable to a causal treatment effect of the home environment, while the remaining 8 percentage points reflect negative selection of less productive workers into remote roles. Remote workers also experienced measurable quality deterioration and a 50% lower promotion rate, echoing Bloom et al. (2015)'s finding across a different industry and context. This treatment-versus-selection decomposition is directly relevant to the present thesis: if workers who most prefer WFH are also those for whom it is most productive, a voluntary regime would generate a different workforce composition than a mandate. Kurdy et al. (2023) corroborate, using survey data from UAE workers during the pandemic, that workload, job satisfaction, WLB, and social support jointly predict remote productivity, while job level has no moderating effect.

Barrero et al. (2023) provide a macro-level perspective on these trends. Full days worked from home rose from 7% in 2019 to 28% by mid-2023, driven by what the authors characterise as a mass social experiment that permanently altered productivity perceptions. Their review of the evidence concludes that fully remote work is associated with some productivity declines, while hybrid arrangements typically maintain or improve performance. The pandemic also redirected substantial innovation toward remote-collaboration technologies and moderated nominal wage growth across sectors, suggesting that WFH has reshaped the returns to complementary inputs as well.

On wages and work burden, the evidence points to substantial heterogeneity by gender and parenthood. Oettinger (2011) documents that home-based employment in the United States nearly doubled between 1980 and 2000 while the associated wage penalty fell from approximately 30 log points to near zero, a trend driven primarily by advances in information technology rather than changes in workforce composition. Within Germany, Arntz et al. (2022) use SOEP data from 1997 to 2014 and find that adopting WFH is associated with roughly one additional unpaid hour per week on average, but also with higher reported job satisfaction. Among parents, WFH reduces gender gaps in contractual hours and monthly income: mothers increase contractual hours by approximately four hours per week, while fathers gain higher hourly wages even without changing employers. The finding that mothers realise wage gains mainly when switching firms points to persistent differences in employer perceptions or bargaining power for remote-working mothers, which may also shape whether mothers with WFH preferences are able to realise them within a given employer relationship.

2.5 Preferences and Willingness to Pay for Remote Work

Less attention has been paid to whether workers actually want remote work and how much they value it. Mas and Pallais (2017) provide the most rigorous evidence on this using a discrete choice experiment embedded in a real national call center recruitment drive across 68 US metro areas. Their central finding is that the average worker is willing to forgo approximately 8% of wages in order to work from home rather than in a standard office. The distribution of this willingness to pay is, however, highly skewed: women, particularly those with young children, place a substantially higher value on WFH and on avoiding employer-controlled irregular schedules than men. The authors also find that scheduling flexibility per se (choosing one's own hours) commands little average value, with 40 hours per week at regular times representing a widely shared preference.

These results imply that aggregate WTP (willingness to pay) figures mask heterogeneity that is both large and systematically correlated with gender and family structure. Combining this with the evidence from Alon et al. (2020) and Lyttelton et al. (2023) on the gendered distribution of domestic labour, the picture that emerges is one where stated preferences for WFH may partly reflect constrained choices: workers who bear the largest domestic burden may value WFH most precisely because it allows them to manage that burden, not because remote work is inherently preferable. Whether this implies a progressive or regressive distributional consequence of WFH expansion depends on whether these same workers are currently able to realise their preferences under a voluntary, employer-driven regime. This is the empirical question this thesis addresses.

2.6 Gap and Contribution

Little of the literature reviewed above directly measures who wants to work from home. The teleworkability literature classifies structural feasibility at the occupation level but largely leaves individual demand unaddressed. The social inequality literature treats realized WFH rates as equivalent to access, making it impossible to distinguish a worker who faces a genuine barrier from one who does not want the arrangement. The well-being and productivity literatures study outcomes conditional on being in a WFH arrangement, but largely leave unaddressed the demand that precedes it. The preferences literature comes closest: Mas and Pallais (2017) show that workers are willing to forgo around 8% of wages for WFH access, with valuations concentrated among women with young children. Their evidence, however, comes from a discrete choice experiment in a single US recruitment context and leaves unanswered whether the workers who value WFH most are actually able to realize it.

This paper exploits an underutilized feature of the SOEP: a self-reported measure of stated WFH preference observed alongside a realized WFH outcome for the same individuals across multiple years. This design allows both stages to be studied separately and a preference-realization gap to be defined for different demographic and socioeconomic groups. Workers who want WFH but cannot obtain it face a genuine constraint; workers who do not want it and do not have it face none. Conflating the two provides an incomplete picture about the distribution of labour market disadvantage and the likely incidence of policies designed to expand remote work access.

3. Data

This study uses data from the German Socio-Economic Panel (Goebel 2024), a nationally representative longitudinal household survey conducted annually since 1984 by the German Institute for Economic Research. The SOEP covers approximately 30,000 individuals in 22,000 households each year, with respondents aged 17 and older, and provides rich individual- and household-level information on income, employment, education, household structure, and demographics. The SOEP is well suited to this analysis because it contains both variables needed to study the preference-realization gap: a direct measure of stated WFH preference and a measure of whether WFH is realized, observed for the same individuals across multiple waves.

The preference variable (`plb0097`) captures responses to the question "If your company allowed you to partially work from home, would you accept this offer?", with three response categories: yes, no, and not possible in my line of work. The third category identifies workers who consider remote work structurally infeasible in their occupation. The realized outcome (`plb0095_v1`) captures whether the worker actually performs their job at home, observed in parallel with the preference variable.

3.1 Sample Description

The analysis covers the 2009–2014 survey waves, in which both variables are consistently available. The sample is restricted to full-time employees (`pgemplst` = 1), excluding part-time workers and the self-employed, for whom the institutional context governing WFH access differs substantially. Workers reporting remote work as structurally infeasible are excluded, so that estimated associations reflect heterogeneity in demand rather than occupational feasibility. The dataset is an unbalanced panel, with workers observed for varying numbers of waves depending on survey participation and employment status. After applying all sample restrictions and dropping observations with missing values on any covariate, the preference-stage sample comprises 7,539 person-year observations. The realization-stage sample is further restricted to workers who stated a positive preference for WFH, yielding 4,975 person-year observations.

Controls include age, sex, direct and indirect migration background, education (ISCED-97), household post-government income (entered as natural logarithm), number of children in the household, housing density (square metres per household member), annual work hours, whether the worker regularly works overtime, and work satisfaction. Industry sector is included via NACE classifications grouped into 15 categories, with Agriculture, Forestry and Fishing as the reference, to account for sectoral differences in remote work prevalence and norms.

Table 2: Descriptive Statistics: Preference and Realization Samples (2009–2014)

Variable	Type	Label	Definition/Scale	Preference Stage			Realization Stage		
				Mean	SD	Within SD	Mean	SD	Within SD
Primary									
`plb0097`	TV	Willing to WFH	0=No, 1=Yes, 3=Not possible	0:34% / 1:66%	–	—	1:100%	–	—
`plb0095_v1`	TV	Actually WFH	0=No, 1=Yes	0:71% / 1:29%	–	—	0:64% / 1:36%	–	—
Demographic									
`age`	TV	Age	years	42.83	9.74	1.32	42.89	9.69	1.29
`sex`	TC	Gender	1=M, 2=F	0:66% / 1:34%	–	—	0:65% / 1:35%	–	—
`has_partner`	TV	Has Partner in HH	0=No, 1=Yes	0:31% / 1:69%	–	—	0:31% / 1:69%	–	—
`migback`	TC	Migr. Background	1=None, 2=Direct, 3=Indirect	1:87% / 2:9% / 3:4%	–	—	1:90% / 2:6% / 3:4%	–	—
`has_children`	TV	Has Children in HH	0=No, 1=Yes	0:46% / 1:54%	–	—	0:48% / 1:52%	–	—
Socioeconomic									
`i11102`	TV	HH Post-Gov Income	EUR	50087.02	28603.59	9718.41	52056.37	28532.24	9600.11
`pgisced97`	TV	Educ. (ISCED-97)	1–6	4.49	1.42	0.10	4.64	1.38	0.11
`sqm_per_head`	TV	Living space per person	sqm/person	43.08	22.97	5.71	44.59	23.27	5.74
Labour									
`plh0173`	TV	Work satisfaction	0–10	7.11	1.96	0.73	7.01	1.98	0.68
`e11101`	TV	Annual work hrs	hours	2226.34	541.88	193.88	2248.02	546.57	178.51
`plb0193_h`	TV	Overtime	0/1†	0.82	0.38	—	0.85	0.35	—
`sector`		Industry Sector	1–15*	–	–	—	–	–	—
N (Observations)				7,539			4,975		

Note: TV = time-varying; TC = time-constant. For categorical variables, the 'Mean' column shows the percentage distribution across categories. Within SD is the standard deviation of the within-individual demeaned variable, reported for continuous time-varying regressors only; binary and time-constant variables show —.

† Manually harmonized from multiple survey versions.

* Derived from `p_nace` (NACE Rev.1) and `p_nace2` (NACE Rev.2).

The specification distinguishes between covariates of primary theoretical interest and controls for individual circumstances and job characteristics. The primary variables of interest fall into two groups. Sex, migration background, and age are fixed characteristics the individual cannot change, each plausibly shaping WFH access through employer behaviour and occupational sorting. Partnership and parental status are equally central, as family composition directly governs both the demand for flexible arrangements and the practical feasibility of working from home. The remaining covariates serve as controls for individual circumstances and job characteristics. Each captures a distinct channel through which WFH access may differ across workers. Age captures life-cycle differences in WFH demand and seniority effects on access. Sex is central to the research question: women are disproportionately likely to hold primary caregiving responsibilities, which may simultaneously strengthen demand for WFH and, through occupational sorting or differential employer treatment, constrain access. Both direct and indirect migration backgrounds are included to capture potential structural and informational barriers, as workers with migration backgrounds may be concentrated in occupations with lower WFH availability or face differential treatment in the allocation of flexible arrangements. Education on the ISCED-97 scale and household post-government income, entered as a natural logarithm to address right-skew, jointly proxy for occupational sorting into remote-compatible roles and the bargaining power to negotiate flexible arrangements.

Household and job characteristics round out the specification. The number of children captures caregiving responsibilities, a well-documented driver of WFH demand, while housing density, measured as square metres per household member, proxies for the practical suitability of the home as a workspace. Annual work hours and a dummy for regular overtime control for work intensity and employer expectations around physical presence; regular overtime may alternatively signal bargaining power or job demands that are difficult to perform remotely. Work satisfaction controls for general job quality and the employment relationship, as workers in more supportive environments are more likely to have access to flexible arrangements.

A descriptive overview of all variables is presented in Table 2.

4. Method

Both research questions (what drives WFH preference among full-time employees, and what determines whether that preference is realized) are addressed using the same modelling framework: pooled logit and probit models supplemented by a Mundlak correlated random effects probit. The two stages differ in two respects. First, the sample: the preference stage uses all full-time employees who did not report WFH as structurally impossible in their occupation, while the realization stage is restricted to workers who stated a positive preference for WFH. Second, the outcome variable: the preference stage models `plb0097` (whether the worker would accept a WFH arrangement if offered), while the realization stage models `plb0095_v1` (whether the worker actually works from home). Where particular groups are systematically less likely to realize their stated preference, this is interpreted as evidence of constrained access rather than voluntary non-participation, and constitutes the preference-realization gap that is the central quantity of interest.

4.1 Pooled Logit and Probit

Four pooled binary choice models are estimated: a logit and a probit, each fitted with and without sector fixed effects. The binary nature of the outcome motivates this choice over a linear probability model: logit and probit cast the problem as a latent variable model, $y_{it} = \mathbf{1}[y_{it}^* > 0]$, which bounds predicted probabilities in $(0, 1)$ by construction and allows the marginal effect of each covariate to vary with the level of the linear index, whereas a linear specification imposes a constant marginal effect and places no restriction on predicted probabilities. The estimating equation takes the form

$$\Pr(y_{it} = 1 \mid \mathcal{S}_{it}) = F\left(\mathbf{x}_{it}^\top \boldsymbol{\beta} + \sum_{s=2}^{15} \delta_s \, \mathbf{1}[\text{sector}_{it} = s]\right) \tag{1}$$

where y_{it} is the binary outcome for worker i in period t, $\mathcal{S}_{it}$ denotes the relevant sample restriction for each stage, and $F = \Lambda$ (logistic CDF) for the logit models and $F = \Phi$ (standard normal CDF) for the probit models. The sector sum is included in columns (2) and (4) only, with Agriculture as the reference category. Sector fixed effects account for the fact that workers sort into industries in ways correlated with both their observed characteristics and their WFH outcomes, and that remote work norms and working conditions vary across sectors beyond what individual-level controls capture.

The two link functions differ in tail behaviour: the logistic distribution assigns more probability mass to extreme values than the standard normal, so logit and probit tend to diverge most when predicted probabilities are concentrated near zero or one. When outcomes are not heavily bunched at the boundaries, as expected here, the two specifications yield nearly identical marginal effects, and reporting both provides a transparent check that conclusions do not depend on the choice of distributional assumption.

Because $\Pr(y_{it} = 1) = F(\mathbf{x}_{it}^{\top}\boldsymbol{\beta})$ is nonlinear, the coefficient β_k does not directly correspond to a change in the outcome probability. Differentiating with respect to x_k yields the individual-specific marginal effect

$$\frac{\partial \Pr(y_{it} = 1)}{\partial x_k} = f(\mathbf{x}_{it}^{\top}\boldsymbol{\beta})\,\beta_k,$$

where $f = F'$ is the density corresponding to F, which varies with the level of the linear index and therefore differs across individuals. Average marginal effects (AMEs), $\widehat{\mathrm{AME}}_k = n^{-1}\sum_{i,t} f(\mathbf{x}_{it}^{\top}\hat{\boldsymbol{\beta}})\,\hat{\beta}_k$, average these over the sample and are interpretable as the average change in the outcome probability associated with a one-unit increase in x_k. Results are reported in Table 3.

4.2 Mundlak Correlated Random Effects Probit

The pooled models treat each observation as independent and assume that unobserved individual heterogeneity is uncorrelated with the regressors. This assumption is unlikely to hold: stable traits such as home environment quality or general disposition toward workplace flexibility are plausibly correlated with both income and education and with the stated preference for WFH. If left unaddressed, this correlation causes the pooled estimates to attribute to observed characteristics what is partly driven by persistent individual differences. Since SOEP tracks the same individuals across multiple waves, the panel structure can be exploited to better control for this heterogeneity.

A standard fixed effects model is not well-suited here for two reasons. First, including individual fixed effects in a nonlinear model such as probit requires estimating one intercept α_i per individual alongside $\boldsymbol{\beta}$,

$$y_{it} = \mathbf{x}_{it}^{\top}\boldsymbol{\beta} + \alpha_i + u_{it} \tag{2}$$

With large N and short T, the number of nuisance parameters grows with the sample, so each $\hat{\alpha}_i$ is estimated from only a handful of observations and remains imprecise. Unlike in the linear case, this imprecision biases $\hat{\boldsymbol{\beta}}$ in a way that does not vanish as $N \to \infty$ (Neyman and Scott 1948). This incidental parameters problem is specific to non-linear models; a linear probability model with individual fixed effects does not suffer from it and is therefore estimated as a specification check for the realization stage (column 7 of Table 4). However, this does not resolve the second limitation. Second, fixed effects identification relies entirely on within-individual variation over time, which means all time-invariant regressors are absorbed by α_i and drop out of the model. Sex and migration background, two variables central to the research question, are time-invariant and would therefore be unidentified.

A standard random effects model avoids both problems but introduces its own: it requires the individual effect α_i to be uncorrelated with the regressors. This is unlikely to hold here, since stable unobserved traits such as home environment quality or employer type are plausibly correlated with income, education, and family structure simultaneously. Treating them as independent when they are not produces the same kind of omitted variable bias as pooled OLS.

To sidestep these shortcomings, this paper employs a model proposed by Mundlak (1978). Rather than treating α_i as independent of the regressors, the individual effect is modelled as a linear function of each individual's time-mean covariates,

$$\alpha_i = \bar{\mathbf{x}}_i^\top \boldsymbol{\pi} + w_i, \qquad \bar{\mathbf{x}}_i = T_i^{-1} \sum_{t=1}^{T_i} \mathbf{x}_{it}, \tag{3}$$

where w_i is a residual that is uncorrelated with $\bar{\mathbf{x}}_i$ by construction. Mundlak (1978, p. 72) shows that $\boldsymbol{\pi} = \mathbf{0}$ corresponds exactly to the standard random effects assumption, making it a testable restriction rather than a maintained one.

Substituting Equation 3 back into the base model yields the augmented specification:

$$y_{it} = \mathbf{x}_{it}^\top \boldsymbol{\beta} + \bar{\mathbf{x}}_i^\top \boldsymbol{\pi} + w_i + u_{it} \tag{4}$$

Mundlak's central result is that the estimator of $\boldsymbol{\beta}$ from this augmented model is numerically identical to the within (fixed effects) estimator, meaning that including the time-means as controls fully accounts for the correlation between α_i and $\mathbf{x}_{it}$. Time-means are computed only for time-varying covariates; for time-constant variables such as sex and migration background, the individual mean equals the variable itself, so including both would introduce perfect collinearity and these variables enter in levels only.

For a binary outcome, the model is cast as a latent variable problem:

$$y_{it}^* = \mathbf{x}_{it}^\top \boldsymbol{\beta} + \bar{\mathbf{x}}_i^\top \boldsymbol{\pi} + w_i + \varepsilon_{it}, \qquad y_{it} = \mathbf{1}[y_{it}^* > 0], \qquad \varepsilon_{it} \overset{iid}{\sim} N(0,1) \tag{5}$$

Assuming $w_i \sim N(0, \omega^2)$, the composite error $e_{it} = w_i + \varepsilon_{it}$ is normally distributed, so the model reduces to a pooled probit on the augmented design matrix $(\mathbf{X}, \bar{\mathbf{X}})$, estimated by maximum likelihood. This normality argument is why the Mundlak device extends cleanly to probit but not to logit, where a normal and a logistic error do not combine into a logistic. Since observations within the same individual share w_i, standard errors are clustered at the individual level. Average marginal effects are reported rather than raw coefficients. The estimating equation is therefore

$$\Pr(y_{it} = 1 \mid \mathcal{S}_{it}) = \Phi\left(\mathbf{x}_{it}^\top \boldsymbol{\beta} + \bar{\mathbf{x}}_i^\top \boldsymbol{\pi} + \sum_{s=2}^{15} \delta_s \, \mathbf{1}[\text{sector}_{it} = s]\right) \tag{6}$$

where $\boldsymbol{\pi} = \mathbf{0}$ is a testable restriction corresponding to the standard random effects assumption. To verify empirically that the Mundlak correction achieves its intended purpose, a linear probability model with individual fixed effects is estimated alongside the CRE probit for the realization stage (column 7 of Table 4). Since individual fixed effects absorb all time-constant variation, sex and migration background cannot be identified in that specification and are omitted from that column.

4.3 Preference and Realization Stages

The full modelling framework, comprising the pooled specifications in Equation 1 and the Mundlak correlated random effects probit in Equation 6, is applied separately to each of the two research questions. Across all specifications, F denotes the relevant link function, and only the outcome variable and estimation sample differ between stages.

The preference stage estimates the probability that a worker would accept WFH if offered (`plb0097`), denoted PTWH_{it} (possibility to work from home). The sample is restricted to full-time employees for whom WFH is not structurally infeasible, a restriction that defines who is observed but is not part of the estimand itself:

$$\Pr(\text{PTWH}_{it} = 1) = F\big(\mathbf{x}_{it}^\top \boldsymbol{\beta}\big) \tag{7}$$

The realization stage estimates the probability of actually working from home (`plb0095_v1`), estimated on the subsample of workers who stated a positive WFH preference ($\text{PTWH}_{it} = 1$):

$$\Pr(\text{WFH}_{it} = 1 \mid \text{PTWH}_{it} = 1) = F\big(\mathbf{x}_{it}^\top \boldsymbol{\beta}\big) \tag{8}$$

The Mundlak time-means $\bar{\mathbf{x}}_i$ are recomputed on the respective estimation sample for each stage. In the realization stage, $\bar{\mathbf{x}}_i$ is therefore averaged only over waves in which $\text{PTWH}_{it} = 1$, making it a conditional mean. Since WFH preference is primarily driven by job characteristics and individual disposition, both slow-moving, this conditional mean is expected to closely approximate the individual's unconditional background characteristics.

The preference-realization gap is identified by comparing AMEs across demographic groups within the realization stage; a group with a lower AME is less likely to obtain

WFH despite having stated a preference, which is interpreted as constrained access rather than voluntary non-participation. This is a within-Stage-2 comparison across groups, not a comparison of aggregate outcomes between Stage 1 and Stage 2. The interpretation assumes that, conditional on observed covariates, selection into $\text{PTWH}_{it} = 1$ is broadly comparable across demographic groups; systematic differences in who states a preference could otherwise confound the estimated gaps.

4.4 Identification Limitations

A central challenge in this analysis is ensuring that the estimated effects are not driven by unobserved factors. While the Mundlak correction accounts for stable individual traits, it cannot capture changes that occur over time alongside the variables I observe. For example, if a worker's bargaining power or specific job responsibilities change at the same time as their income or household structure, the results might still be affected by omitted variable bias. This is a particular concern for the realization stage, where things like a worker's relationship with their manager or specific employer-side constraints are likely important but unobserved.

Furthermore, the model assumes that the included controls capture the main reasons why people sort into different work arrangements. If unobserved traits influence both the desire for and the ability to work from home in a way that changes over time, the estimated gaps might reflect these selection effects rather than purely structural barriers. For these reasons, the findings are best interpreted as clear associations and patterns rather than definitive causal proof.

Table A1 in Appendix A summarises all twelve specifications estimated across both stages.

5. Results

The results are presented in two subsections corresponding to the two research questions. The first subsection addresses the determinants of stated WFH preferences among feasible workers, drawing on the pooled and correlated random effects models estimated on the full analysis sample. The second subsection examines which characteristics predict whether a stated preference is realized, estimated on the subsample of workers who expressed a positive preference for WFH.

Figure 1 provides a descriptive overview of WFH demand and realization rates across demographic groups before turning to the regression analysis. The figure illustrates the two patterns that motivate the two-stage design: women have higher demand than men but lower realization, while workers with a direct migration background have markedly lower demand than workers without a migration background.

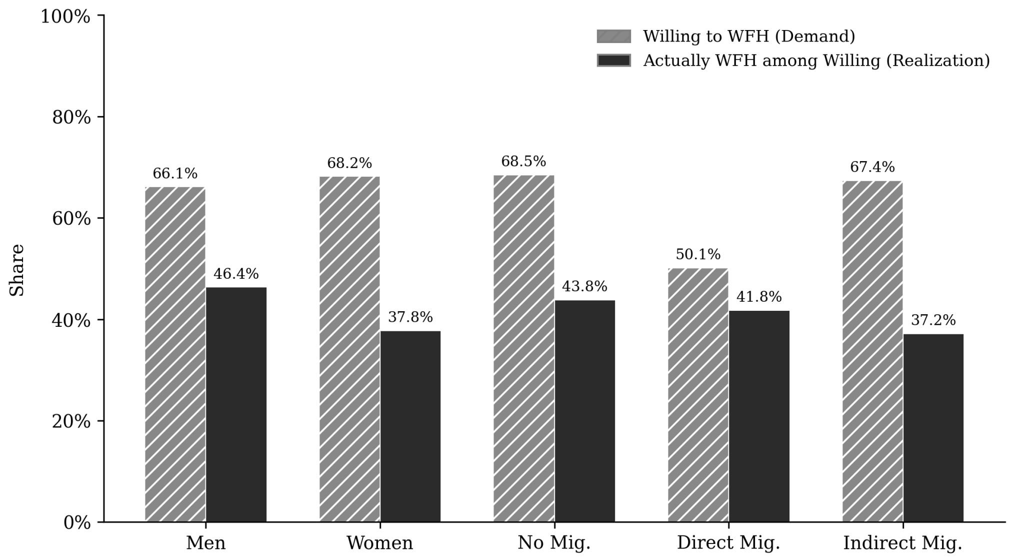

Figure 1: WFH demand and realization rates by demographic group. Demand is the share of workers willing to WFH; realization is the share actually working from home among those willing.

5.1 Stated Preferences for Remote Work

Table 3 reports average marginal effects from the pooled logit, pooled probit, and correlated random effects probit, each estimated with and without sector fixed effects. The logit and probit estimates are near-identical throughout, so the discussion focuses on the probit results in columns (3) and (4).

In the pooled probit without sector fixed effects (column 3), female workers are about 3 percentage points more likely to want WFH (AME = 0.030, $p < 0.05$). Workers with a direct migration background are about 14 percentage points less likely to want WFH

(AME $= -0.141$, $p < 0.01$); indirect migration background is not significant in any specification. Education and income are both positively associated with WFH preferences: a one-unit increase in ISCED-97 education corresponds to a 3.3 percentage point higher probability, and a one log-point increase in household income to an 8.5 percentage point higher probability, both significant at the 1% level. More living space per household member is also positively associated with WFH preferences (AME $= 0.0014$, $p < 0.01$). Working overtime is associated with about a 10 percentage point higher probability of wanting WFH ($p < 0.01$), while higher work satisfaction is associated with lower WFH demand (AME $= -0.018$, $p < 0.01$). Age has a small negative effect (-0.0020, $p < 0.01$), and having children in the household is not significant.

Adding sector fixed effects in columns (2) and (4) changes little. The education AME falls from 0.033 to 0.029 and the income AME from 0.085 to 0.078, indicating that part of the education and income associations run through sectoral sorting rather than individual-level demand. Other coefficients are stable.

The Mundlak means are jointly significant in both CRE specifications ($p < 0.001$), indicating that the pooled models are biased by unobserved individual heterogeneity and that the CRE correction is warranted.

In the CRE specification, education, living space, and overtime all lose significance; household income weakens to the 10% level. The mechanism differs across variables. Education is near-time-constant (within-individual SD $= 0.10$; Table 2): its Mundlak mean captures nearly the same information as the level variable, and the CRE cannot separately identify the two, so the coefficient is absorbed into the mean. For living space and overtime, the attenuation reflects between-individual selection: workers with more living space or regular overtime tend to have stable unobserved characteristics that predict WFH preferences, and once those are absorbed by the Mundlak correction the coefficients disappear. Three variables remain significant across all six models. Female workers show a stable 3 percentage point preference advantage. Workers with a direct migration background show a stable 14 percentage point lower probability of wanting WFH. Work satisfaction retains a small negative effect (AME $= -0.014$, $p < 0.05$), meaning less satisfied workers are more likely to want WFH, and this holds within individuals over time. Workers who report higher satisfaction likely have stronger reasons to be at the workplace, reducing demand for remote arrangements.

Having a partner in the household is not significant in any pooled specification (AME ≈ -0.006, $p > 0.1$ across columns 1–4), yet the CRE probit yields a large and significant negative effect (AME $= -0.104$, $p < 0.01$). This pattern indicates that the cross-sectional correlation between partnership status and WFH preferences is confounded by stable individual characteristics; once those are absorbed by the Mundlak correction, a within-person signal becomes more discernible. Workers who gain a partner reduce their stated demand for WFH by around 10 percentage points, which may reflect a preference for separating the domestic and work domains once cohabitation introduces competing claims on home space and routine.

The absence of a significant children effect is consistent with Mas and Pallais (2017), who find that WFH demand is driven primarily by being female rather than by parenthood per se. Their evidence that mothers with young children value WFH more than other women is captured in the present model through the female indicator rather than a standalone children variable (Mas and Pallais 2017, p. 3752).

That household income weakens but does not fully disappear in the CRE suggests its pooled association partly reflects stable individual differences, while a genuine within-person relationship also exists: when the same worker's income rises over time, their stated WFH preference tends to rise modestly with it.

The McFadden R^2 of approximately 0.05 across all specifications indicates a modest fit. McFadden (1977) notes that values between 0.2 and 0.4 represent excellent fit for discrete choice models, so the present models explain only a limited share of the variation in WFH preferences. This reflects genuine heterogeneity in individual demand that observed characteristics do not fully capture.

Finally, I address the possibility that stated WFH preferences are endogenous to current WFH status. If workers who already work from home develop a stronger preference for it as they adapt their lifestyle and domestic routines, the observed associations might reflect this reinforcement rather than underlying demand. To address this concern for reverse-causality, Table A3 in Appendix A re-estimates the models on the subsample of workers who do not yet work from home (`plb0095_v1` $= 0$). The results are qualitatively identical: the female preference advantage remains significant and slightly larger than in the full sample (AME ≈ 0.045, $p < 0.01$), and the migration background penalty persists. Education and income also retain their positive associations in the pooled models. This strengthens the interpretation that the primary determinants of WFH demand are not merely artifacts of current work arrangements, but represent genuine differences in stated demand among those yet to attain flexible arrangements.

Table 3: Willingness to Work from Home: Pooled and Correlated Random Effects Models

	Logit		**Probit**		**CRE Probit**	
	(1)	(2)	(3)	(4)	(5)	(6)
Demographic Characteristics						
Age	-0.0021^{***}	-0.0019^{***}	-0.0020^{***}	-0.0019^{***}	-0.0028	-0.0027
	(0.0007)	(0.0007)	(0.0007)	(0.0007)	(0.0034)	(0.0034)
Female	0.0288^{**}	0.0277^{**}	0.0294^{**}	0.0286^{**}	0.0305^{**}	0.0298^{**}
	(0.0135)	(0.0138)	(0.0135)	(0.0138)	(0.0135)	(0.0139)
Direct Migr. Background	-0.1371^{***}	-0.1326^{***}	-0.1412^{***}	-0.1367^{***}	-0.1383^{***}	-0.1340^{***}
	(0.0199)	(0.0200)	(0.0207)	(0.0207)	(0.0207)	(0.0207)
Indirect Migr. Background	0.0007	0.0031	0.0001	0.0030	0.0014	0.0044
	(0.0318)	(0.0314)	(0.0318)	(0.0314)	(0.0318)	(0.0314)
Has Partner in HH	-0.0075	-0.0057	-0.0061	-0.0040	-0.1044^{***}	-0.1037^{***}
	(0.0146)	(0.0146)	(0.0146)	(0.0145)	(0.0391)	(0.0391)
Socioeconomic Determinants						
Education (ISCED-97)	0.0322^{***}	0.0283^{***}	0.0325^{***}	0.0287^{***}	-0.0219	-0.0205
	(0.0045)	(0.0047)	(0.0045)	(0.0047)	(0.0326)	(0.0334)
HH Post-Gov. Income†	0.0875^{***}	0.0807^{***}	0.0852^{***}	0.0782^{***}	0.0676^{*}	0.0667^{*}
	(0.0147)	(0.0148)	(0.0144)	(0.0146)	(0.0402)	(0.0403)
Has Children in HH	-0.0005	0.0067	-0.0018	0.0052	0.0547	0.0586
	(0.0143)	(0.0144)	(0.0142)	(0.0144)	(0.0397)	(0.0399)
Living Space (sqm/person)	0.0015^{***}	0.0014^{***}	0.0014^{***}	0.0013^{***}	-0.0009	-0.0008
	(0.0004)	(0.0004)	(0.0004)	(0.0004)	(0.0009)	(0.0009)
Labour Determinants						
Annual Work Hours (per 100h)	0.0008	0.0011	0.0009	0.0012	-0.0016	-0.0017
	(0.0011)	(0.0011)	(0.0011)	(0.0011)	(0.0020)	(0.0020)
Works Overtime	0.0982^{***}	0.1028^{***}	0.0993^{***}	0.1039^{***}	0.0231	0.0260
	(0.0138)	(0.0138)	(0.0142)	(0.0142)	(0.0247)	(0.0246)
Work Satisfaction	-0.0177^{***}	-0.0175^{***}	-0.0177^{***}	-0.0176^{***}	-0.0139^{**}	-0.0137^{**}
	(0.0029)	(0.0029)	(0.0029)	(0.0029)	(0.0055)	(0.0055)
Sector FE	No	Yes	No	Yes	No	Yes
N (all models)	7,539					
McFadden R^2 / Within R^2	0.0458	0.0542	0.0453	0.0538	0.0474	0.0559
Log-Likelihood	-4610.8	-4570.2	-4612.9	-4572.1	-4602.8	-4562.0
Mundlak means (p)	—	—	—	—	< 0.001	< 0.001

Note: Average marginal effects (AME) reported. Cluster-robust delta-method standard errors in parentheses, clustered at the individual level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. † Variable entered as natural logarithm. CRE Probit includes individual time-means of time-varying regressors (Mundlak correction). Sector fixed effects: NACE sectors 1–15 (reference: Agriculture).

5.2 Realization of Stated Preferences

Table 4 reports average marginal effects from the pooled logit, pooled probit, correlated random effects probit, and a linear probability model with individual fixed effects (FE LPM), each estimated with and without sector fixed effects where applicable. Since logit and probit again yield near-identical estimates across all specifications, the probit results in columns (3) and (4) serve as the primary reference for the discussion below.

In the pooled probit without sector fixed effects (column 3), female workers who stated a positive preference for WFH are about 5.6 percentage points less likely to actually work from home (AME $= -0.056$, $p < 0.01$). Education and income are strongly positively associated with WFH realization: a one-unit increase in ISCED-97 education corresponds to a 6.8 percentage point higher probability, and a one log-point increase in household income to a 16.5 percentage point higher probability, both significant at the 1% level. Having children in the household is associated with a 9.0 percentage point lower probability of realizing WFH ($p < 0.01$). More living space per household member is positively associated with WFH realization (AME $= 0.0014$, $p < 0.01$), and each additional 100 annual work hours with a 0.9 percentage point higher probability ($p < 0.01$). Work satisfaction is positively associated with WFH realization (AME $= 0.014$, $p < 0.01$), reversing the direction observed in the preference stage; workers with high satisfaction appear more likely to have employers who maintain high job quality and simultaneously permit remote work. Age is small and positively significant (AME $= 0.0015$, $p < 0.05$). Migration background, partnership status, and overtime are not significant in the pooled specification.

Adding sector fixed effects in columns (2) and (4) reduces the education AME from 0.068 to 0.051, indicating that part of the education premium in WFH access reflects sorting into sectors with higher remote-work prevalence. The contribution of sector to model fit is notably larger here than in the preference stage: the probit McFadden R^2 rises from 0.129 to 0.183 when sector dummies are included. This is consistent with WFH access being more directly shaped by industry and occupational norms than WFH demand.

The Mundlak means are jointly significant in both CRE specifications ($p < 0.001$), confirming that unobserved individual heterogeneity is correlated with the regressors and that the pooled estimates are biased. To validate the CRE approach, column (7) reports an FE LPM, which avoids the incidental parameters bias inherent to nonlinear fixed-effects models and imposes no distributional assumption on the error. Because individual fixed effects absorb all time-constant variation, sex and migration background cannot be identified in the FE LPM; the column serves as a within-estimator check on the time-varying regressors only.

In the CRE specifications, education, household income, children, living space, and work satisfaction all lose significance. For education, the attenuation is primarily attributable to near-zero within-individual variation: in the realization sample, the within-individual standard deviation of ISCED-97 is only 0.11 (Table 2), leaving the Mundlak mean and the level variable nearly collinear, so the coefficient is absorbed into the mean. For income, children, and living space, the pattern reflects between-individual selection: workers with higher incomes, without children, or with more living space sort into jobs and circumstances more conducive to WFH, and once stable individual heterogeneity is absorbed by the Mundlak correction, within-person variation in these characteristics does not predict changes in WFH realization. The strong pooled associations for education and income are consistent with the projections of Bonacini et al. (2021) and the descriptive evidence in Yasenov (2020), who document that WFH access is concentrated among higher-educated, higher-income workers; the CRE results qualify this picture differently for each variable: income's disappearance reflects between-individual sorting into WFH-compatible roles, while education's attenuation is attributable to near-zero within-individual variation rather than an identified sorting mechanism. The children result is also consistent with the preference stage: as Mas and Pallais (2017) show, WFH demand is driven primarily by being female rather than by parenthood, and the absence of a within-person children effect in the realization stage suggests that parental status shapes job selection rather than employer decisions within comparable positions. Work satisfaction similarly loses significance in the CRE, suggesting its pooled association captures stable employer-quality differences rather than a genuine within-person dynamic.

Three variables remain significant across the CRE and FE LPM specifications. Annual work hours retain a positive association (AME ≈ 0.008, $p < 0.01$), consistent with workers logging more hours having stronger incentives to negotiate flexible arrangements. Overtime, not significant in any pooled specification, emerges positive and significant in the CRE and FE LPM (AME ≈ 0.068, $p < 0.05$), indicating that the overtime-WFH association reflects within-person transitions toward remote arrangements rather than cross-sectional sorting. Age is substantially amplified in the CRE relative to the pooled probit (AME $= 0.021$ vs 0.0015): the cross-sectional age pattern partially reflects stable individual characteristics correlated with both age and WFH access, and once those are absorbed, accumulating seniority predicts meaningfully higher WFH realization within individuals.

The most robust finding across all specifications is the female access penalty. Female workers who want WFH are between 5.6 and 6.7 percentage points less likely to work from home, with the estimate stable across pooled logit, pooled probit, and CRE probit specifications. This contrasts sharply with the preference stage, where female workers were more likely to state a positive WFH preference. The female preference advantage in Stage 1 is consistent with Mas and Pallais (2017), who find that women's willingness to pay for WFH substantially exceeds men's; the access penalty observed here indicates that this demand premium has not translated into equivalent access. A related pattern is documented by Arntz et al. (2022), who find that German mothers realize WFH-related earnings gains primarily through firm switches rather than within existing employment relationships, pointing to persistent within-employer resistance to flexible arrangements for women.

Migration background, a strong negative predictor of WFH demand in the preference stage, is not significant in any realization-stage specification.

The McFadden R^2 ranges from approximately 0.13 without sector fixed effects to 0.19 with sector fixed effects, substantially higher than the approximately 0.05 observed in the preference stage. The consistency between FE LPM and CRE probit estimates on time-varying covariates nonetheless confirms that the CRE results are not an artefact of the distributional assumptions underlying the probit.

As a robustness check, Table A4 in Appendix A re-estimates the same specifications on the full preference-stage sample, with stated WFH preference (`plb0097`) included as a covariate rather than used as a sample restriction. The results are broadly consistent with the main analysis: the female penalty remains negative and significant across all six specifications, and the Mundlak correction again renders education, income, children, and work satisfaction insignificant. The female AME is somewhat smaller at approximately 3.6 percentage points compared to 5.6 percentage points in the main analysis, which is intuitive if the preference-restricted sample disproportionately captures women who face particularly high barriers to obtaining WFH. That the gap persists and remains significant on the broader sample suggests it is not simply an artifact of conditioning on stated preference. One notable difference is that direct migration background becomes marginally significant (10%) once sector fixed effects are included, a pattern that does not emerge in the main analysis, hinting at a broader access deficit that may be partially obscured when restricting to workers who actively want WFH.

Even with these checks, it is possible that some factors are still missing from the model. Unobserved job details, like the specific nature of daily tasks or how much autonomy someone has, could still influence both demographic characteristics and the likelihood of working from home. The results point to clear and systematic differences in WFH access that persist even after accounting for many individual and job-level characteristics, but it is likely that unobserved factors also play a role in how these arrangements are allocated.

Table 4: Realization of WFH Preferences: Access and Utilization among Willing Workers

	Logit		Probit		CRE Probit		FE LPM
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Demographic Characteristics							
Age	0.0016**	0.0014*	0.0015**	0.0014*	0.0205***	0.0213***	0.0211***
	(0.0008)	(0.0007)	(0.0008)	(0.0007)	(0.0040)	(0.0040)	(0.0041)
Female	−0.0552***	−0.0683***	−0.0556***	−0.0669***	−0.0561***	−0.0672***	—
	(0.0155)	(0.0152)	(0.0155)	(0.0151)	(0.0155)	(0.0152)	
Direct Migr. Background	−0.0024	0.0013	−0.0011	0.0007	−0.0009	0.0003	—
	(0.0269)	(0.0251)	(0.0272)	(0.0255)	(0.0273)	(0.0256)	
Indirect Migr. Background	−0.0096	−0.0080	−0.0032	0.0000	−0.0042	−0.0016	—
	(0.0356)	(0.0341)	(0.0352)	(0.0340)	(0.0351)	(0.0340)	
Has Partner in HH	0.0146	0.0113	0.0162	0.0133	−0.0567	−0.0564	−0.0586
	(0.0166)	(0.0158)	(0.0166)	(0.0158)	(0.0500)	(0.0494)	(0.0484)
Socioeconomic Determinants							
Education (ISCED-97)	0.0676***	0.0500***	0.0683***	0.0505***	−0.0024	−0.0204	−0.0066
	(0.0051)	(0.0052)	(0.0051)	(0.0052)	(0.0636)	(0.0674)	(0.0628)
HH Post-Gov. Income†	0.1706***	0.1832***	0.1648***	0.1783***	−0.0049	0.0009	−0.0033
	(0.0173)	(0.0167)	(0.0172)	(0.0167)	(0.0568)	(0.0542)	(0.0520)
Has Children in HH	−0.0895***	−0.0701***	−0.0901***	−0.0708***	0.0362	0.0429	0.0467
	(0.0160)	(0.0153)	(0.0159)	(0.0153)	(0.0487)	(0.0498)	(0.0493)
Living Space (sqm/person)	0.0013***	0.0015***	0.0014***	0.0016***	0.0003	0.0005	0.0004
	(0.0003)	(0.0003)	(0.0003)	(0.0003)	(0.0010)	(0.0010)	(0.0011)
Labour Determinants							
Annual Work Hours (per 100h)	0.0099***	0.0088***	0.0093***	0.0085***	0.0076***	0.0072***	0.0069***
	(0.0014)	(0.0013)	(0.0014)	(0.0013)	(0.0027)	(0.0027)	(0.0026)
Works Overtime	0.0152	0.0252	0.0163	0.0250	0.0678**	0.0749**	0.0647**
	(0.0192)	(0.0185)	(0.0192)	(0.0185)	(0.0312)	(0.0331)	(0.0306)
Work Satisfaction	0.0138***	0.0120***	0.0136***	0.0118***	−0.0018	−0.0010	−0.0019
	(0.0035)	(0.0033)	(0.0034)	(0.0033)	(0.0062)	(0.0062)	(0.0060)
Sector FE	No	Yes	No	Yes	No	Yes	No
N (all models)	4,975						
McFadden R^2 / Within R^2	0.1298	0.1838	0.1286	0.1825	0.1327	0.1868	0.0541
Log-Likelihood	-2838.5	-2662.6	-2842.6	-2666.9	-2829.2	-2652.6	—
Mundlak means (p)	—	—	—	—	< 0.001	< 0.001	—

Note: Average marginal effects (AME) reported. Cluster-robust delta-method standard errors in parentheses, clustered at the individual level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. † Variable entered as natural logarithm. CRE Probit includes individual time-means of time-varying regressors (Mundlak correction). Sector fixed effects: NACE sectors 1–15 (reference: Agriculture).

6. Discussion

The introduction motivated this analysis on three grounds: that the welfare-relevant inequality is not who works from home but who wants to and cannot; that access-side interventions such as a statutory WFH right are only meaningful where the preference-realization gap is large; and that a systematic failure to realize stated preferences points to employer-side bargaining power over the terms of the employment relationship. The findings are relevant to all three, but the picture differs across demographic groups. For female workers, the gap is on the access side and holds after within-individual estimation, making it the clearest evidence of a genuine access barrier in the data. For workers with a direct migration background, the gap is entirely on the demand side, with no evidence of differential access once a preference is stated. For income, the panel structure shows that the realization-stage association largely reflects occupational sorting rather than employer-side barriers within comparable positions, while a within-person component remains in the preference stage. For education, the CRE result reflects a data limitation rather than a substantive finding: near-zero within-individual variation in attainment prevents the panel from identifying a within-person effect.

The female finding is the paper's central result. The preference advantage in Stage 1 and the access penalty in Stage 2 together show that women express stronger demand for WFH but are less likely to obtain it, with the access penalty holding after the Mundlak correction. Framed in terms of the compensating differentials argument outlined in the introduction, women place a higher value on WFH than men yet cannot obtain it at the same rate. If workers would willingly accept lower wages in exchange for WFH and still cannot obtain it, the market is not delivering the outcome the theory predicts, and the access penalty effectively functions as an implicit wage penalty.

Several mechanisms could explain this gap. Within teleworkable industries, women may be concentrated in client-facing or supervisory roles where physical presence carries stronger informal norms, a form of within-sector occupational segregation that sector fixed effects cannot capture. Employer reluctance may also reflect assumptions about reduced commitment or domestic distractions. A further possibility is that employers factor in the career cost associated with remote work: Bloom et al. (2015) document that home workers are promoted at half the rate of equally productive office colleagues, and employers who treat this as a productivity signal may selectively extend WFH to groups perceived as more committed to the workplace. As noted in the results, German mothers tend to achieve WFH-related earnings gains primarily through firm switches rather than within existing employers, consistent with persistent within-employer resistance to flexible arrangements for women.

The migration background results reveal a different type of inequality. Workers with a direct migration background are about 14 percentage points less likely to want WFH in Stage 1, yet face no access disadvantage conditional on having stated a preference. The inequality is entirely on the demand side. Several explanations are plausible: workers earlier in their careers may place greater value on workplace visibility for performance signalling, and migrants disproportionately occupy such positions; communication barriers in a second language may raise the perceived cost of remote coordination; cultural norms around workplace presence may also differ across origin countries. Whatever the reason, migrants are not being denied WFH; they are simply less likely to want it. This distinction matters for interpretation: the finding is not evidence of employer discrimination but of lower stated demand, which may itself reflect rational responses to labour market conditions. The absence of any significant effect for indirect migration background is consistent with second-generation workers being more assimilated to the host-country labour market.

The pooled associations between education, income, and WFH outcomes are consistent with prior work documenting that WFH is concentrated among highly educated, high-income workers. The CRE results, however, suggest a different picture. In the realization stage, both education and income lose significance: for education, this reflects near-zero within-individual variation in attainment rather than an identified sorting mechanism; for income, it is consistent with between-individual sorting into WFH-compatible roles. In the preference stage, income retains marginal significance (AME $= 0.068$, $p < 0.1$), indicating a genuine within-person component alongside the sorting effect. The gap in WFH uptake between high- and low-educated workers cannot be attributed to a within-person mechanism in this panel: within-individual variation in educational attainment is near-zero, so the CRE cannot identify an education effect separately from stable individual traits. The same applies to children: the negative pooled association in Stage 2 reflects selection into roles with lower WFH availability rather than within-person changes in WFH access as family circumstances change.

The reversal in the work satisfaction association between stages is also worth discussing. Dissatisfied workers are more likely to want WFH, consistent with treating remote arrangements as an escape from an unpleasant workplace environment, yet satisfied workers are more likely to realize it. Satisfaction likely captures stable features of the employment relationship, such as employer quality, job autonomy, and trust, that also predict whether an employer will grant flexible arrangements. The pattern is consistent with Bellmann and Hübler (2020)'s finding that formally contracted arrangements yield higher satisfaction than informal ones, and that the termination of existing WFH access significantly worsens work-life balance, suggesting that WFH access and employment relationship quality may correlate. Neither effect holds after the Mundlak correction, confirming that both reflect stable individual-employer matching rather than causal within-person dynamics.

6.1 Policy Implications

Three policy-relevant conclusions follow from the results.

The female access gap provides the clearest empirical basis for access-side intervention. The gap is large, holds after within-individual estimation, and reflects a demand premium that the market has not accommodated. A statutory right to WFH for workers in feasible occupations, of the kind debated in several European legislatures (SPD 2020; Milders 2022; Citizens Information 2024; UNECE 2020), would on these results disproportionately benefit women who currently want the arrangement and cannot obtain it. The compensating differentials framing strengthens this case: if employers are denying an arrangement that workers would willingly accept at lower wages, requiring equal access is a correction of a market failure rather than a purely redistributive transfer.

The migration background results caution against treating all WFH gaps as access-side problems. For workers with a direct migration background, the constraint is on the demand side, and access-side legislation would not address a lower stated preference for WFH. Designing policy around aggregate WFH uptake statistics would misdiagnose the problem for this group. Understanding why stated demand is lower among migrants, whether through signalling concerns, communication costs, or cultural norms, is necessary before identifying the right policy response, if any.

The education and income results imply that WFH expansion would not straightforwardly reduce socioeconomic inequality in labour market outcomes. Because the education and income gap in WFH uptake reflects sorting into occupations rather than unequal access within comparable jobs, a statutory right to WFH would primarily benefit workers already in suitable roles. The more fundamental inequality operates upstream, at the point of occupational sorting, and access policy alone cannot easily fix it.

6.2 Limitations

The analysis is subject to several limitations. The data cover 2009 to 2014, predating the structural shift in employer norms brought about by the COVID-19 pandemic. Remote work expanded dramatically from 2020 onward, and the magnitude and pattern of preference-realization gaps may have changed as employers accumulated direct evidence that WFH is feasible across a wider range of roles. The magnitude of the gaps documented here may therefore differ from what a post-pandemic analysis would find, though the structural patterns are likely to remain informative.

The preference measure is a stated rather than revealed preference: workers report whether they would accept WFH if offered, not whether they have actively sought it. Workers who report a positive preference may differ in unobserved intensity of demand, and the hypothetical framing may be susceptible to social desirability bias. The robustness check on workers who do not (yet) work from home, reported in Table A3, addresses reverse causality, but the hypothetical nature of the measure remains a limitation of the design.

Finally, key employer-side determinants of WFH realization are unobserved. Firm-level policies, manager discretion, and the specific nature of daily tasks all plausibly affect whether a stated preference translates into a realized arrangement. The gender result also has an important limitation: since sex does not change over time, it cannot be identified from within-person variation and is instead estimated from comparisons between men and women. This means that stable unobserved differences between men and women cannot be ruled out. One important example is that women may sort into different types of jobs within the same sector in ways that affect WFH availability, and sector fixed effects cannot capture this. These are the most plausible sources of remaining bias in the gender result.

7. Conclusion

This paper used panel data from the German Socio-Economic Panel (SOEP) for 2009–2014 to study both who wants to work from home and who actually obtains it, among full-time employees in teleworkable occupations. By separating stated preferences from realized outcomes, the analysis moves beyond prior work that treats raw WFH uptake rates as a direct measure of access. The results show that the two questions have distinct answers across demographic groups. Women are more likely to want WFH but significantly less likely to obtain it, a gap that persists after controlling for observed characteristics and individual heterogeneity. Workers with a direct migration background show the opposite pattern: a large demand deficit but no additional access disadvantage conditional on having stated a preference. For income, the realization-stage association largely reflects occupational sorting rather than within-job access barriers. Education's pooled association cannot be given a within-person interpretation from this panel, as variation in attainment over time is insufficient for identification.

These findings suggest that WFH inequality is not a single phenomenon and cannot be addressed by a single policy instrument. The female access gap points to employer-side constraints that access-side policy could target, while the migration demand gap raises different questions about preferences and labour market integration. Future research would benefit from post-pandemic data, where the large expansion of WFH may have narrowed some gaps while widening others, and from employer-level data that would allow the role of firm policies and manager discretion to be estimated more directly.

References

Alon, Titan et al. (Apr. 2020). *The Impact of COVID-19 on Gender Equality*. w26947. Cambridge, MA: National Bureau of Economic Research, w26947. DOI: 10.3386/w26947. URL: http://www.nber.org/papers/w26947.pdf (visited on 04/28/2026).

Arntz, Melanie, Sarra Ben Yahmed, and Francesco Berlingieri (June 1, 2022). "Working from Home, Hours Worked and Wages: Heterogeneity by Gender and Parenthood". In: *Labour Economics* 76, p. 102169. ISSN: 0927-5371. DOI: 10.1016/j.labeco.2022.102169. URL: https://www.sciencedirect.com/science/article/pii/S0927537122000604 (visited on 05/01/2026).

Bamieh, Omar and Lennart Ziegler (June 1, 2022). "Are Remote Work Options the New Standard? Evidence from Vacancy Postings during the COVID-19 Crisis". In: *Labour Economics* 76, p. 102179. ISSN: 0927-5371. DOI: 10.1016/j.labeco.2022.102179. URL: https://www.sciencedirect.com/science/article/pii/S0927537122000707 (visited on 05/02/2026).

Barrero, José María, Nicholas Bloom, and Steven J. Davis (Nov. 2023). "The Evolution of Work from Home". In: *Journal of Economic Perspectives* 37.4, pp. 23–50. ISSN: 0895-3309. DOI: 10.1257/jep.37.4.23. URL: https://www.aeaweb.org/articles?id=10.1257/jep.37.4.23 (visited on 04/29/2026).

Beland, Louis-Philippe, Abel Brodeur, and Taylor Wright (Mar. 15, 2023). "The Short-Term Economic Consequences of COVID-19: Exposure to Disease, Remote Work and Government Response". In: *PLOS ONE* 18.3, e0270341. ISSN: 1932-6203. DOI: 10.1371/journal.pone.0270341. URL: https://journals.plos.org/plosone/article?id=10.1371/journal.pone.0270341 (visited on 04/28/2026).

Bellmann, Lutz and Olaf Hübler (2020). *Job Satisfaction and Work-Life Balance: Differences between Homework and Work at the Workplace of the Company*. Working Paper 13504. IZA Discussion Papers. URL: https://www.econstor.eu/handle/10419/223946 (visited on 04/27/2026).

Berg, Janine, Bonnet Florence, and Soares Sergei (2020). *Working from Home: A Potential Measure for Mitigating the COVID-19 Pandemic | OIT/Cinterfor*. URL: https://www.cinterfor.org/node/7784 (visited on 05/02/2026).

Bloom, Nicholas et al. (Feb. 1, 2015). "Does Working from Home Work? Evidence from a Chinese Experiment *". In: *The Quarterly Journal of Economics* 130.1, pp. 165–218. ISSN: 0033-5533. DOI: 10.1093/qje/qju032. URL: https://doi.org/10.1093/qje/qju032 (visited on 05/01/2026).

Bonacini, Luca, Giovanni Gallo, and Sergio Scicchitano (Jan. 1, 2021). "Working from Home and Income Inequality: Risks of a 'New Normal' with COVID-19". In: *Journal of Population Economics* 34.1, pp. 303–360. ISSN: 1432-1475. DOI: 10.1007/s00148-

020-00800-7. URL: https://doi.org/10.1007/s00148-020-00800-7 (visited on 04/28/2026).

Cavapozzi, Danilo (Aug. 1, 2022). "The Causal Impact of Remote Working on Depression during the First Wave of the COVID-19 Pandemic". In: *French Stata Users' Group Meetings 2022* 10. URL: https://ideas.repec.org//p/boc/fsug22/10.html (visited on 04/28/2026).

Choudhury, Prithwiraj (Raj), Cirrus Foroughi, and Barbara Larson (2021). "Work-from-Anywhere: The Productivity Effects of Geographic Flexibility". In: *Strategic Management Journal* 42.4, pp. 655–683. ISSN: 1097-0266. DOI: 10.1002/smj.3251. URL: https://onlinelibrary.wiley.com/doi/abs/10.1002/smj.3251 (visited on 05/01/2026).

Citizens Information (2024). *Right to Request Remote Working*. URL: https://www.citizensinformation.ie/en/employment/employment-rights-and-conditions/contracts-of-employment/right-to-request-remote-working/ (visited on 05/03/2026).

Destasis (2023). *Employed Persons Working from Home*. Federal Statistical Office. URL: https://www.destatis.de/EN/Themes/Labour/Labour-Market/Quality-Employment/Dimension3/3_11_homeoffice.html (visited on 05/03/2026).

Dingel, Jonathan I. and Brent Neiman (Sept. 1, 2020). "How Many Jobs Can Be Done at Home?" In: *Journal of Public Economics* 189, p. 104235. ISSN: 0047-2727. DOI: 10.1016/j.jpubeco.2020.104235. URL: https://www.sciencedirect.com/science/article/pii/S0047272720300992 (visited on 04/28/2026).

Emanuel, Natalia and Emma Harrington (Oct. 2024). "Working Remotely? Selection, Treatment, and the Market for Remote Work". In: *American Economic Journal: Applied Economics* 16.4, pp. 528–559. ISSN: 1945-7782. DOI: 10.1257/app.20230376. URL: https://www.aeaweb.org/articles?id=10.1257/app.20230376 (visited on 05/01/2026).

Felstead, Alan and Darja Reuschke (2020). "HOMEWORKING IN THE UK: BEFORE AND DURING THE 2020 LOCKDOWN". In.

García-Salirrosas, Elizabeth Emperatriz et al. (Nov. 2023). "Job Satisfaction in Remote Work: The Role of Positive Spillover from Work to Family and Work–Life Balance". In: *Behavioral Sciences* 13.11, p. 916. ISSN: 2076-328X. DOI: 10.3390/bs13110916. URL: https://www.mdpi.com/2076-328X/13/11/916 (visited on 04/27/2026).

Goebel, Jan (2024). *German Socio-Economic Panel Study (SOEP-Core)*. DOI: 10.5684/soep.core.v41o.

Gottlieb, Charles et al. (Apr. 2021). "Working from Home in Developing Countries". In: *European Economic Review* 133, p. 103679. ISSN: 00142921. DOI: 10.1016/j.euroecorev.2021.103679. URL: https://linkinghub.elsevier.com/retrieve/pii/S0014292121000325 (visited on 05/02/2026).

Gueguen, Guillaume and Claudia Senik (2023). "Adopting Telework: The Causal Impact of Working from Home on Subjective Well-Being". In: *British Journal of Industrial Relations* 61.4, pp. 832–868. ISSN: 1467-8543. DOI: 10.1111/bjir.12761. URL: https://onlinelibrary.wiley.com/doi/abs/10.1111/bjir.12761 (visited on 04/28/2026).

Holgersen, Henning, Zhiyang Jia, and Simen Svenkerud (Dec. 1, 2021). "Who and How Many Can Work from Home? Evidence from Task Descriptions". In: *Journal for Labour Market Research* 55, p. 4. DOI: 10.1186/s12651-021-00287-z.

Kurdy, Daniela M., Husam-Aldin N. Al-Malkawi, and Shahid Rizwan (2023). "(PDF) The Impact of Remote Working on Employee Productivity during COVID-19 in the UAE: The Moderating Role of Job Level". In: *ResearchGate*. DOI: 10.1108/JBSED-09-2022-0104. URL: https://www.researchgate.net/publication/367269352_The_impact_of_remote_working_on_employee_productivity_during_COVID-19_in_the_UAE_the_moderating_role_of_job_level (visited on 05/01/2026).

Lyttelton, Thomas, Emma Zang, and Kelly Musick (2023). "Parents' Work Arrangements and Gendered Time Use during the COVID-19 Pandemic". In: *Journal of Marriage and Family* 85.2, pp. 657–673. ISSN: 1741-3737. DOI: 10.1111/jomf.12897. URL: https://onlinelibrary.wiley.com/doi/abs/10.1111/jomf.12897 (visited on 04/28/2026).

Mas, Alexandre and Amanda Pallais (Dec. 2017). "Valuing Alternative Work Arrangements". In: *American Economic Review* 107.12, pp. 3722–3759. ISSN: 0002-8282. DOI: 10.1257/aer.20161500. URL: https://www.aeaweb.org/articles?id=10.1257/aer.20161500 (visited on 05/01/2026).

McFadden, Daniel (Nov. 1, 1977). "Quantitative Methods for Analyzing Travel Behaviour of Individuals: Some Recent Developments". In: *Cowles Foundation Discussion Papers*. URL: https://elischolar.library.yale.edu/cowles-discussion-paper-series/707.

Milders, Lydia (Aug. 23, 2022). *Netherlands: The Legislative Proposal Working Where You Want Act*. L&E Global. URL: https://leglobal.law/2022/08/23/netherlands-the-legislative-proposal-working-where-you-want-act/ (visited on 05/03/2026).

Montenovo, Laura et al. (May 2020). *Determinants of Disparities in Covid-19 Job Losses*. w27132. Cambridge, MA: National Bureau of Economic Research, w27132. DOI: 10.3386/w27132. URL: http://www.nber.org/papers/w27132.pdf (visited on 04/28/2026).

Mundlak, Yair (1978). "On the Pooling of Time Series and Cross Section Data". In: *Econometrica* 46.1, pp. 69–85. ISSN: 0012-9682. DOI: 10.2307/1913646. JSTOR: 1913646. URL: https://www.jstor.org/stable/1913646 (visited on 05/04/2026).

Neyman, J. and Elizabeth L. Scott (1948). "Consistent Estimates Based on Partially Consistent Observations". In: *Econometrica* 16.1, pp. 1–32. ISSN: 0012-9682. DOI: 10.

2307/1914288. JSTOR: 1914288. URL: https://www.jstor.org/stable/1914288 (visited on 05/05/2026).

Oettinger, Gerald S. (2011). "The Incidence and Wage Consequences of Home-Based Work in the United States, 1980-2000". In: *The Journal of Human Resources* 46.2, pp. 237–260. ISSN: 0022-166X. JSTOR: 41304819. URL: https://www.jstor.org/stable/41304819 (visited on 05/01/2026).

Perelman, Julian et al. (Jan. 2021). "Does Working at Home Compromise Mental Health? A Study on European Mature Adults in COVID Times". In: *Journal of Occupational Health* 63.1, e12299. ISSN: 1348-9585. DOI: 10.1002/1348-9585.12299. PMID: 34894172.

Ploszaj, Heloise et al. (Apr. 7, 2025). "Understanding the Associations between "Work from Home", Job Satisfaction, Work-Life Balance, Stress, and Gender in an Organizational Context of Remote Work". In: *Discover Psychology* 5.1, p. 24. ISSN: 2731-4537. DOI: 10.1007/s44202-025-00342-7. URL: https://doi.org/10.1007/s44202-025-00342-7 (visited on 04/27/2026).

Sanchez, Daniel Garrote et al. (2020). "Who on Earth Can Work from Home?" In.

Schifano, Sonia (2023). "Well-Being and Working from Home during COVID-19 | Request PDF". In: *ResearchGate*. DOI: 10.1108/ITP-01-2021-0033. URL: https://www.researchgate.net/publication/356841158_Well-being_and_working_from_home_during_COVID-19 (visited on 04/28/2026).

Senik, Claudia et al. (Feb. 2, 2024). "Teleworking and Life Satisfaction during COVID-19: The Importance of Family Structure". In: *Journal of Population Economics* 37.1, p. 8. ISSN: 1432-1475. DOI: 10.1007/s00148-024-00979-z. URL: https://doi.org/10.1007/s00148-024-00979-z (visited on 04/27/2026).

Sostero, Matteo et al. (2020). "Teleworkability and the COVID-19 Crisis: A New Digital Divide?" In: URL: https://publications.jrc.ec.europa.eu/repository/handle/JRC121193 (visited on 05/02/2026).

Sostero, Matteo et al. (Dec. 17, 2023). "Teleworkability and the COVID-19 Crisis: Potential and Actual Prevalence of Remote Work across Europe". In: *IZA Journal of Labor Policy* 13. DOI: 10.2478/izajolp-2023-0006.

SPD (2020). *Arbeit, die zum Leben passt*. Sozialdemokratische Partei Deutschlands (SPD). URL: https://www.spd.de/aktuelles/detail/news/arbeit-die-zum-leben-passt/05/10/2020 (visited on 04/30/2026).

UNECE (Sept. 22, 2020). *Real Decreto-ley 28/2020, de 22 de Septiembre, de Trabajo a Distancia*. URL: https://ageing-policies.unece.org/browse-policy/802 (visited on 05/03/2026).

Xiao, Yijing et al. (Mar. 1, 2021). "Impacts of Working From Home During COVID-19 Pandemic on Physical and Mental Well-Being of Office Workstation Users". In: *Journal* *of Occupational and Environmental Medicine* 63.3, pp. 181–190. ISSN: 1536-5948. DOI: 10.1097/JOM.0000000000002097. PMID: 33234875.

Yasenov, Vasil (Apr. 2020). “Who Can Work from Home?” In: *IZA Discussion Papers* 13197. URL: https://ideas.repec.org//p/iza/izadps/dp13197.html (visited on 04/28/2026).

A. Tables

Table A1: Model Specifications and Sampling Logic

Model ID	Estimator	Sector FE	Functional Form (Latent Variable y_{it}^*)
(1), (3)	Logit, Probit	No	$\mathbf{x}_{it}^\top\beta + \epsilon_{it}$
(2), (4)	Logit, Probit	Yes	$\mathbf{x}_{it}^\top\beta + \gamma_s + \epsilon_{it}$
(5)	CRE Probit	No	$\mathbf{x}_{it}^\top\beta + \bar{\mathbf{x}}_i^\top\pi + \epsilon_{it}$
(6)	CRE Probit	Yes	$\mathbf{x}_{it}^\top\beta + \bar{\mathbf{x}}_i^\top\pi + \gamma_s + \epsilon_{it}$
Application to Research Questions			
Stage 1 (RQ1)	Outcome: Preference ($y_{it} = 1$). Sample: Full-time, feasible ($N = 7,539$).		
Stage 2 (RQ2)	Outcome: Realization ($y_{it} = 1$). Sample: Full-time, willing ($N = 4,975$).		

Note: $\mathbf{x}_{it}$ is the vector of all control variables. $\bar{\mathbf{x}}_i$ denotes individual time-means of time-varying variables (Mundlak correction). γ_s represents sector fixed effects (NACE 1–15).

Table A2: NACE Industry Sector Classification

#	Sector	NACE Rev. 1 codes	NACE Rev. 2 codes
1	Agriculture, Forestry & Fishing	1–2, 5	1–3
2	Mining & Quarrying	10–14	5–9
3	Manufacturing	15–37	10–33
4	Utilities & Waste	40–41	35–39
5	Construction	45	41–43
6	Trade & Hospitality	50–52, 55	45–47, 55–56
7	Transport & Logistics	60–63	49–53
8	Information & Communication	64, 72	58–63
9	Finance & Insurance	65–67	64–66
10	Professional & Business Services	70–71, 73–74	68–75, 77–82
11	Public Administration	75	84
12	Education	80	85
13	Health & Social Work	85	86–88
14	Other Services	90–93, 95–97	90–98
15	Extra-territorial	99	99

Note: Rev.1 and Rev.2 codes are mapped separately before coalescing. Rev.1 takes priority where both are observed.

Table A3: Willingness to Work from Home (Restricted to non-WFH workers)

	Logit		**Probit**		**CRE Probit**	
	(1)	(2)	(3)	(4)	(5)	(6)
Demographic Characteristics						
Age	-0.0028^{***}	-0.0027^{***}	-0.0028^{***}	-0.0027^{***}	-0.0060	-0.0061
	(0.0008)	(0.0008)	(0.0008)	(0.0008)	(0.0048)	(0.0048)
Female	0.0477^{***}	0.0443^{***}	0.0486^{***}	0.0451^{***}	0.0491^{***}	0.0455^{***}
	(0.0166)	(0.0171)	(0.0166)	(0.0171)	(0.0167)	(0.0172)
Direct Migr. Background	-0.1719^{***}	-0.1654^{***}	-0.1752^{***}	-0.1686^{***}	-0.1728^{***}	-0.1662^{***}
	(0.0247)	(0.0246)	(0.0251)	(0.0250)	(0.0251)	(0.0250)
Indirect Migr. Background	0.0093	0.0094	0.0098	0.0098	0.0113	0.0114
	(0.0378)	(0.0376)	(0.0378)	(0.0376)	(0.0378)	(0.0377)
Has Partner in HH	-0.0169	-0.0146	-0.0160	-0.0135	-0.0746	-0.0728
	(0.0177)	(0.0176)	(0.0176)	(0.0175)	(0.0467)	(0.0470)
Socioeconomic Determinants						
Education (ISCED-97)	0.0237^{***}	0.0231^{***}	0.0238^{***}	0.0233^{***}	-0.0552	-0.0457
	(0.0057)	(0.0058)	(0.0057)	(0.0058)	(0.0372)	(0.0376)
HH Post-Gov. Income†	0.0583^{***}	0.0464^{***}	0.0568^{***}	0.0453^{***}	0.0557	0.0546
	(0.0180)	(0.0179)	(0.0175)	(0.0175)	(0.0507)	(0.0511)
Has Children in HH	0.0291	0.0338^{*}	0.0282	0.0329^{*}	-0.0498	-0.0505
	(0.0179)	(0.0180)	(0.0178)	(0.0179)	(0.0548)	(0.0546)
Living Space (sqm/person)	0.0018^{***}	0.0018^{***}	0.0018^{***}	0.0017^{***}	-0.0000	0.0000
	(0.0005)	(0.0005)	(0.0004)	(0.0004)	(0.0014)	(0.0014)
Labour Determinants						
Annual Work Hours (per 100h)	-0.0006	-0.0003	-0.0006	-0.0003	0.0007	0.0006
	(0.0014)	(0.0014)	(0.0014)	(0.0014)	(0.0024)	(0.0023)
Works Overtime	0.1227^{***}	0.1265^{***}	0.1236^{***}	0.1272^{***}	0.0009	0.0040
	(0.0167)	(0.0167)	(0.0170)	(0.0170)	(0.0292)	(0.0293)
Work Satisfaction	-0.0228^{***}	-0.0222^{***}	-0.0226^{***}	-0.0221^{***}	-0.0122^{*}	-0.0122^{*}
	(0.0035)	(0.0035)	(0.0035)	(0.0035)	(0.0066)	(0.0066)
Sector FE	No	Yes	No	Yes	No	Yes
N (all models)	5,356					
McFadden R^2 / Within R^2	0.0425	0.0505	0.0424	0.0504	0.0446	0.0525
Log-Likelihood	-3469.6	-3440.8	-3470.0	-3441.2	-3462.2	-3433.5
Mundlak means (p)	—	—	—	—	< 0.001	< 0.001

Note: Average marginal effects (AME) reported. Cluster-robust delta-method standard errors in parentheses, clustered at the individual level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. † Variable entered as natural logarithm. CRE Probit includes individual time-means of time-varying regressors (Mundlak correction). Sector fixed effects: NACE sectors 1–15 (reference: Agriculture).

Table A4: WFH Realization: Full Eligible Sample Robustness (Preference Unrestricted, plb0097 Controlled)

	Logit		**Probit**		**CRE Probit**	
	(1)	(2)	(3)	(4)	(5)	(6)
WFH Preference						
WFH Preference (stated)	0.1822***	0.1674***	0.1779***	0.1640***	0.1767***	0.1627***
	(0.0114)	(0.0112)	(0.0111)	(0.0108)	(0.0111)	(0.0108)
Demographic Characteristics						
Age	0.0012**	0.0012**	0.0012**	0.0013**	0.0173***	0.0183***
	(0.0006)	(0.0005)	(0.0006)	(0.0005)	(0.0029)	(0.0029)
Female	−0.0373***	−0.0493***	−0.0357***	−0.0473***	−0.0357***	−0.0471***
	(0.0119)	(0.0118)	(0.0118)	(0.0117)	(0.0119)	(0.0118)
Direct Migr. Background	−0.0309	−0.0295*	−0.0329*	−0.0335*	−0.0327*	−0.0333*
	(0.0191)	(0.0178)	(0.0188)	(0.0176)	(0.0189)	(0.0177)
Indirect Migr. Background	−0.0065	−0.0019	−0.0021	0.0058	−0.0033	0.0042
	(0.0272)	(0.0261)	(0.0270)	(0.0261)	(0.0269)	(0.0261)
Has Partner in HH	0.0061	0.0036	0.0057	0.0031	−0.0653*	−0.0699*
	(0.0125)	(0.0121)	(0.0125)	(0.0120)	(0.0380)	(0.0369)
Socioeconomic Determinants						
Education (ISCED-97)	0.0603***	0.0465***	0.0610***	0.0470***	0.0134	−0.0006
	(0.0039)	(0.0039)	(0.0038)	(0.0039)	(0.0469)	(0.0480)
HH Post-Gov. Income†	0.1277***	0.1371***	0.1225***	0.1325***	0.0267	0.0222
	(0.0128)	(0.0125)	(0.0127)	(0.0125)	(0.0395)	(0.0380)
Has Children in HH	−0.0760***	−0.0625***	−0.0766***	−0.0630***	0.0340	0.0382
	(0.0120)	(0.0116)	(0.0120)	(0.0115)	(0.0396)	(0.0395)
Living Space (sqm/person)	0.0012***	0.0013***	0.0012***	0.0013***	0.0003	0.0004
	(0.0003)	(0.0002)	(0.0003)	(0.0002)	(0.0007)	(0.0007)
Labour Determinants						
Annual Work Hours (per 100h)	0.0089***	0.0080***	0.0084***	0.0077***	0.0063***	0.0060***
	(0.0011)	(0.0010)	(0.0011)	(0.0010)	(0.0021)	(0.0021)
Works Overtime	0.0360**	0.0479***	0.0382***	0.0487***	0.0573**	0.0669***
	(0.0141)	(0.0137)	(0.0139)	(0.0136)	(0.0248)	(0.0254)
Work Satisfaction	0.0106***	0.0087***	0.0100***	0.0081***	−0.0027	−0.0024
	(0.0027)	(0.0026)	(0.0027)	(0.0026)	(0.0050)	(0.0050)
Sector FE	No	Yes	No	Yes	No	Yes
N (all models)	7,535					
McFadden R^2 / Within R^2	0.1694	0.2171	0.1683	0.2159	0.1722	0.2201
Log-Likelihood	-3763.9	-3548.0	-3768.9	-3553.4	-3751.4	-3534.5
Mundlak means (p)	—	—	—	—	< 0.001	< 0.001

Note: Average marginal effects (AME) reported. Cluster-robust delta-method standard errors in parentheses, clustered at the individual level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. † Variable entered as natural logarithm. CRE Probit includes individual time-means of time-varying regressors (Mundlak correction). Sector fixed effects: NACE sectors 1–15 (reference: Agriculture).