The Distributional Consequences of Public Pay Caps

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Abstract

We study the distributional consequences of a common post-financial crisis policy: public-sector pay caps. In the UK, caps limited nominal pay growth to 1% from 2010 to 2017 for public workers earning above £21,000 (15% of employees). Using counterfactual distribution methods, we find larger *public* wage losses at P90 than P50. The caps reduced the *overall* P90-P50 gap by 2.3pp by 2019, fully accounting for the observed decline in inequality, while widening gender and regional wage gaps by 0.9pp and 0.4pp. The contemporaneous decline in public-sector employment approximately doubled regional and gender effects to 1.9pp (17.5%) and 0.7pp (9.9%).

Key words: Austerity, Public Sector, Pay Caps, Inequality

JEL: J24, J31, J45

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1 Introduction

The public sector plays a central role in developed economies by providing essential services with positive externalities—such as education, healthcare, and policing. Across the OECD, it accounts for approximately 18% of total employment (OECD, 2025), making competitive public wage policy a key driver of productivity and output (Britton & Propper, 2016; Burgess et al., 2017; Dal Bó et al., 2013; Nickell & Quintini, 2002; Propper & Van Reenen, 2010). However, in response to the global financial crisis, governments of over 25 IMF high-income countries implemented public pay restraint policies—including wage freezes and caps—as part of broader fiscal consolidation efforts (Ortiz et al., 2015). Many of these policies remained in place considerably after the financial crisis, and the long-run effects have been associated with widespread industrial action (Cribb & O'Brien, 2024).[1]

This paper studies the unintended distributional consequences of such policies across places and sexes as well as the overall distribution. While existing work has focused on consequences within the public sector, we leverage a novel extension of the DiNardo et al., 1996 counterfactual technique to consider the impact of public sector wage regimes across the entire wage distribution.

We address these phenomena by examining a well-defined UK policy: the public sector pay freeze from 2010 to 2017, which applied to all public employees earning above £21,000 (approximately $35,000 in 2026 USD). Introduced as part of a broader austerity programme, the policy aimed to reduce government spending from 45% to 38% of national income and eliminate a 10% budget deficit (Emmerson, 2017). Its duration, scope, and clear design make the UK an ideal case study, with approximately half of all public sector workers—over 15% of total employees—being affected. Despite binding based on annual pay, we centre our analysis around hourly wages to abstract away from intensive labour supply adjustments and focus on a metric closer to marginal productivity. Our results are robust to changing outcomes.

[1] See Di Porto et al. (2023) for an example of a salary cap that was only repealed in 2025.

Over the policy period, real hourly wage inequality in the UK—as measured by the P90–P10 ratio—fell by over 9%, reversing a 30% rise between 1980 and 2010 (Bell et al., 2022; Giupponi et al., 2024). The decline was driven by falling wages at the top and rising wages at the bottom. Most explanations emphasise the latter—such as the National Living Wage (Dube, 2019; Giupponi & Machin, 2026; Giupponi et al., 2024) and rising demand for low-skilled work (Giupponi & Machin, 2024)—but cannot account for the decline at the 90th percentile. We revisit this contraction by focusing on the P90–P50 gap and assess the extent to which public pay restraint explains the fall in top-end inequality.

In a first stage, we document a 4.5pp reduction in the *public sector* hourly wage P90–P50 ratio between 2010 and 2019. Throughout, we focus on 2019 to demonstrate that, although the pay caps ended in 2017, their scarring effects persist unless pay growth subsequently accelerates.[2] Over the same period, the conditional public sector wage premium – the average wage differential between observationally equivalent public and private sector workers – fell from 13.7% to 8.6% driven disproportionately by the top end of the wage distribution, reducing the monetary return to working in the public sector. We then simulate counterfactual wage growth along the public sector distribution, assuming that the return to observable characteristics in the public sector mirrored those in the private sector using an extension of the DiNardo et al. (1996) decomposition method, which we refer to as the *relative price effect*. We then use the institutional timing and threshold structure of the pay caps to assess whether the resulting relative-return effect is plausibly attributable to public pay restraint. Under this counterfactual scenario, the public sector P90–P50 ratio would have increased by 2.1pp, providing convincing evidence that the pay caps did indeed bite.

Treating private sector wages as observed, we extend our analysis to show that the *overall* P90–P50 ratio in the economy would have been 2.3pp higher in 2019 under the same counterfactual scenario. The pronounced effect on overall inequality arises because public sector workers are disproportionately concentrated at the upper end of the overall wage

[2]We stop short of examining our counterfactuals beyond 2020 to abstract from the effects of the COVID-19 period.

distribution. This underscores the importance of public sector pay policies in shaping broader income inequality trends during this period.

With nearly 65% of public sector workers being women, the pay caps disproportionately affected female wages. Our counterfactual analysis indicates that changes to public sector pay policy led to an 8.3% increase in the gender pay gap in real terms, relative to the counterfactual scenario. Similarly, the public sector accounts for a smaller share of employment in the South of England than in the rest of the UK, and its share outside the South has grown since 2010. As a result, workers outside the South were more exposed to public pay austerity, leading to a 6.0% increase in the South's wage premium over the rest of the UK and exacerbating already stark regional disparities.

Turning to changes in composition, we show that shifting worker characteristics further suppressed wage growth along the distribution in the public sector by 0.4pp, although these effects were uniform along the distribution. While a similar average effect is true in the private sector, the underlying compositional shifts occurred for different reasons. In the public sector, shifts in the age and sex distribution contributed negatively to wage growth, but this was partially offset by a movement toward full time work. In contrast, wage growth in the private sector decreased because of a movement away from full time work, and the negative effects of shifting age and sex composition played a much smaller role.

These changes in composition were largely driven by increased outflow from the public sector. After 2010, the public sector retention rate—measured as the share of public sector workers who remain in the sector the following year—fell by 8.1pp, reversing the pattern observed in the previous seven years, during which public sector retention consistently exceeded that of the private sector. Among workers exiting the public sector, the majority transitioned directly into private sector employment rather than into unemployment.

Our final counterfactual accounts for the 7.0pp reduction in the public sector employment share from 30.8% to 23.8% between 2010 and 2019 resulting from increased net outflows. Because observationally equivalent public sector workers earned a real wage premium, the declining public sector share disproportionately reduced wages for women and workers outside the South of England, who were more heavily represented in public employment. Along these margins, the public sector share effect contributed as much to rising inequality as the pay caps themselves.

Accounting for shifts in relative price, composition and the public share together, our counterfactuals suggest that the gender pay gap would have been 1.9pp (17.5%) lower and the regional gap would have been 0.7pp (9.9%) lower. The picture is different for the P90–P50 ratio. Because the public sector wage distribution is compressed relative to the private sector, the counterfactual increase in the public sector share leaves the P90 relatively unchanged while substantially increasing the P50. The decline in the public sector share therefore increased P90–P50 inequality between 2010 and 2018. This effect was large enough to offset 70% of the reduction in inequality attributable to the relative price effect.

We find relatively little evidence that the pay caps led to selective changes in inflow or outflow from the public sector. We take two steps to evaluate that phenomenon. First, we examine inflow and outflow rates along the wage distribution and show that they are relatively uniform throughout. Second, we replicate our characteristic counterfactual but include inflow and outflow indicators as observable characteristics. We find that changes in inflow and outflow patterns made little contribution to shifts in the wage distribution, and that any estimated effects were relatively uniform across the distribution. Our results instead point toward the role of contemporaneous austerity policies other than the pay caps—such as departmental budget cuts resulting in selective job destruction (Facchetti, 2024; Villa, 2024)—in shaping the composition of the public sector workforce over this period. This null result also strengthens the interpretation of our relative price effect as capturing the distributional consequences of the pay caps.

We contribute to two literatures. The first concerns the determinants of wage inequality, with a focus on the UK. One strand attributes its evolution to market forces: skill-biased technological change widened differentials between high- and low-skilled workers;

Goos and Manning (2007) link the hollowing out of middle-earning work to routine-task substitution; and Bell and Van Reenen (2014) trace the rising top 1% earnings share to pay at the top. A second strand studies institutions: declining union coverage (Gosling & Lemieux, 2004; Machin, 1997) and the National Living Wage, which raised wages at the bottom with spillovers to around the 20th percentile (Dube, 2019; Giupponi et al., 2024).

Across this literature, however, the public sector typically enters as a control rather than as a distinct wage-setting regime. This matters because the mechanisms emphasised above—skill demand and task substitution—primarily operate through market wage-setting, whereas around 28.5% of UK employees work in the public sector, where pay is more commonly shaped by national agreements, administrative rules, and compressed pay structures. Changes originating in public pay policy are therefore liable to be attributed to market forces. The omission extends beyond the UK: canonical institutional accounts likewise do not treat the public sector separately (DiNardo et al., 1996; Lemieux, 2006).

The second literature concerns public-sector wage setting. The state sets pay for a large share of workers through national, compressed pay schedules relative to the private sector, making public pay policy a distinct determinant of the aggregate distribution. Most studies estimate the public–private differential, with two findings. First, the large positive raw public-sector premium significantly shrinks or reverses once workforce composition is accounted for (Borjas, 2002; Dustmann & van Soest, 1998; Postel-Vinay & Turon, 2007). Second, the differential favours low-paid workers and penalises high-paid workers, so public pay is more compressed than private pay (Bargain et al., 2026). This differential has reduced since the financial crisis, internationally (Michael & Christofides, 2020) and in the UK (Murphy et al., 2020). Yet its implications for aggregate inequality remain unmeasured. We take the differential as a starting point rather than a conclusion.

Bradley et al. (2017) provide the closest analysis of how public pay affects movements between sectors. In a Burdett-Mortensen show that cuts to public pay reduce public employment mainly through voluntary reallocation into the private sector, while leaving private-sector wages largely unchanged. Their concern is with worker sorting across sectors; ours is with the effect of public pay policy and subsequent changes in composition on the wage distribution itself.

An additional strand of public-sector wage setting literature studies the uneven incidence of nationally determined public pay across workers and places. National pay structures can generate regional differences in public-sector recruitment and labour allocation (Boeri et al., 2021). In the UK, regulated pay has also been linked to productivity and service outcomes in schools and hospitals (Britton & Propper, 2016; Propper & Van Reenen, 2010). Related work examines women's selection into public employment (Gomes & Kuehn, 2026) and how public-sector institutions shape gender pay gaps within the sector (Biasi & Sarsons, 2022). While previous work considers the within-sector implications, Jones et al. (2018) considers aggregate outcomes: they show that a high public sector share can reduce gender inequality in the UK because of the high fraction of women in high-wage public sector jobs, which suppresses the gender pay gap within the public sector relative to the private sector. The subsequent reduction in the public sector share since the financial crisis therefore coincided with an increase in the gender pay gap by 2015, although the estimates are quantitatively small because gender wage inequalities exist within both sectors. We instead ask how public-sector returns have changed since 2010 and what this implies for the gender pay gap, regional wage differences, and overall inequality.

The paper proceeds as follows: Section 2 describes our linked employer-employee data. Section 3 outlines changes to the public sector wage structure before and after 2010, along with the relevant austerity policies. Section 4 details our counterfactual methodology and presents summary statistics. Section 5 presents our main results on overall inequality, the gender pay gap, and the regional wage divide. Section 6 concludes.

2 Data

The primary dataset for this project is the Annual Survey of Hours and Earnings (ASHE). Our linked employer-employee data span 1997 to 2019 and link workers to their employers at the individual level. The dataset comprises a 1% random sample of employees, selected based on the last two digits of their National Insurance number—the UK equivalent of the US Social Security number. ASHE is the foremost source of earnings data in the UK and serves as the basis for the Office for National Statistics' (ONS) official earnings statistics. It has also been widely used in both economic and policy research.

Individual tax records—collected via His Majesty's Revenue and Customs (HMRC) Pay As You Earn (PAYE) system—are linked to firms registered on the Inter-Departmental Business Register (IDBR). Data are collected directly from employers, which minimises attenuation bias typically associated with self-reported surveys, such as the Labour Force Survey (LFS). Participation is legally mandated, and the response rate is approximately 80%, corresponding to around 180,000 individuals annually.

Despite being a theoretically random sample, certain individuals are not observed. Firms with annual turnover below £85,000 ($115,000) are not required to register with the IDBR, implying that many self-employed sole traders are omitted. Bell et al. (2022) note that these workers account for approximately 15% of total employment. Moreover, workers who do not pay National Insurance—those earning up to approximately £120 ($161) per week—are not registered in the PAYE system and therefore cannot be linked to a firm.[3] This is equivalent to a minimum wage worker employed for 15 hours per week, representing a very small proportion of the population, primarily composed of women in the private sector. We therefore apply the predefined ASHE sampling weights throughout, which aim to align the sample of ASHE respondents with the representative employee population.

A public sector worker is defined as an employee in central or local government, or in public corporations.[4] In the UK, this includes occupations such as teachers and civil servants.

[3] National Insurance is the statutory tax that all workers pay to be eligible for a state pension.

[4] This follows the formal definition adopted by the Office for National Statistics, the UK's official statistical body.

Unlike in the US, the public sector also includes the majority of healthcare professionals, who are employed by the National Health Service (NHS). Across our sample, public workers account for 27.0% of total employment, suggesting that any change to public sector pay has implications for the overall wage distribution.

Our primary outcome variable is gross hourly wage, calculated as weekly earnings (including overtime) divided by the number of hours worked in the reference week.[5] Additional variables include sex, age, and local authority region.

Educational attainment is not observed in ASHE. This is a potential limitation in our setting, where the average educational attainment of public sector workers differs from that of private sector workers. To address this, we replicate parts of our analysis using the Labour Force Survey (LFS) where appropriate. The LFS is a quarterly household survey covering approximately 35000 households with a rotating panel structure, from which we construct annual repeated cross-sections for our analysis. Importantly, educational attainment is recorded for each individual in six broad categories: degree or equivalent, higher education, A-level or equivalent, GCSE or equivalent, other qualifications, and no qualifications. Despite this additional information on educational attainment, we continue to use ASHE as our primary data source to avoid the non-response and reporting biases associated with self-reported survey data.

While the linked employer-employee data provide high-quality earnings information, they do not capture labour market states beyond employment in the public or private sector. To address this, we supplement our analysis with the harmonised Understanding Society and British Household Panel Surveys (UKLS). The UKLS is a self-reported longitudinal dataset covering approximately 40,000 individuals in the UK and is widely used in the macro-labour literature (Bradley et al., 2017; Sepahsalari & Postel-Vinay, 2023).[6] Crucially, it provides information on employment type (e.g. self-employment) as well as on non-employment states—including unemployment, inactivity, and retirement—that are

[5]Wages are deflated by the Consumer Price Index (CPI)—the headline measure of inflation in the UK—indexed to 2019.

[6]The UKLS refers to the combination of two surveys: the British Household Panel Survey (BHPS) and Understanding Society (USoC).

not observable in the data. This allows us to construct a more complete picture of worker transitions and sectoral flows.

3 Public Sector Wages and Composition

Next, we describe trends in real wages and inequality in the public and private sectors (Section 3.1). We focus our distributional analysis on the P90–P50 differential. This captures inequality in the upper half of the wage distribution, where the pay caps were most consequential, while avoiding lower percentiles that are more exposed to changes in the wage floor and to compositional changes associated with unemployment.[7][8] Section 3.2 then explores the demographic composition of the public sector and how it evolved after 2010. Finally, Section 3.3 shows that changes in the relative composition of the public sector were predominantly driven by increased outflows at the top end of the wage distribution.

3.1 Trends in Real Wages

Figure 1 shows wage growth an extended period from 1997 to 2019 in the public and private sectors at the 90th percentile, normalised to 1997. Wages increased at the upper end of the distribution in both sectors between 1997 and 2010, with public sector wage growth closely tracking that of the private sector. In contrast, wages declined across both sectors between 2011 and 2015, as the UK experienced the prolonged effects of the Financial and Eurozone crises. From 2015 onward, real wages began to recover in the private sector, while public sector wages continued to decline in real terms. Between 2010 and 2018, public sector workers at the 90th percentile experienced 8% lower real wage growth than their private sector counterparts.

[7]The National Living Wage was introduced in the UK in 2016 for workers aged 25 and over, with the aim of raising the wage floor to 60% of median earnings by 2020. The policy is estimated to affect 7% of individuals aged 25 or over (Dube, 2019).

[8]Between 2010 and 2018, the unemployment rate fell from 7.9% to 4.1%. Blundell et al. (2014) show that the initial unemployment shock was concentrated among young, low-educated male workers, making the lower percentiles of the observed wage distribution particularly susceptible to changes in employment composition.

We hypothesise that much of the change in the public wage distribution can be attributed to the public sector pay freezes. The policy was introduced in 2010 under Chancellor of the Exchequer George Osborne, who froze the nominal wages of all public sector employees earning above £21,000 based on their full-time equivalent salary, with the goal of saving £3.3 billion per year.[9] Workers below the cap threshold were entitled to a minimum raise of £250 per year.

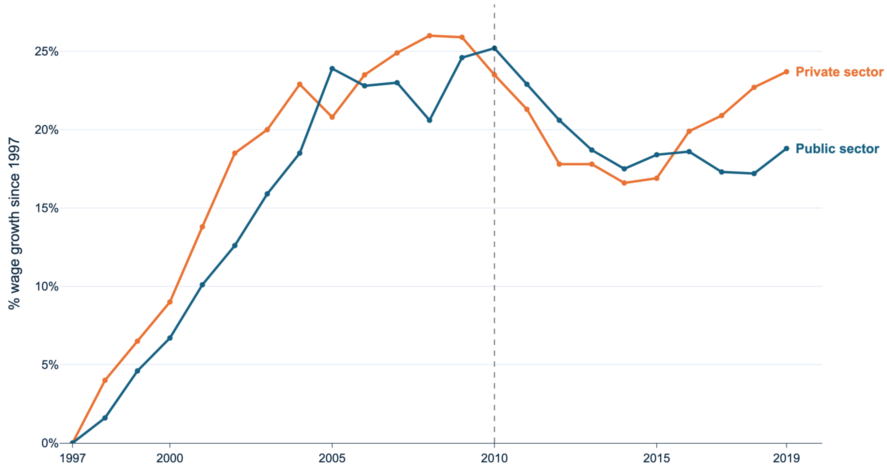


Figure 1: Real P90 Wages Indexed to 1997

Note:. The figure plots the percentage evolution of real hourly wages at the 90th percentile indexed to 1997 = 0. All wages are deflated using the Consumer Price Index (CPI), expressed in 2019 prices. *Source:* ASHE.

The nominal pay freeze remained in effect until 2013, when it was loosened to permit 1% annual wage increases. In 2017, the policy was formally repealed, and all public sector workers became entitled to a minimum 2% nominal pay rise.[10] Over the same period, inflation averaged between 1–3%, resulting in persistent real wage declines for workers subject to the cap.

In our main analysis we assume that, in the absence of the pay caps, wage growth above

[9]The 2011–13 freeze was directly implemented for workforces under UK ministerial control and through the UK pay-review-body system. Local-authority pay was separately negotiated, although the main collective bargaining body for local government employees nonetheless received no national award for three years. Pay policy for devolved public bodies was formally set by the devolved administrations, which generally adopted closely similar restraint).

[10]The policy was repealed as the government neared its target of balancing the budget and faced increasing union pressure due to the widening gap between public and private sector wage growth.

the bite point of the policy would have mirrored that of the private sector. This assumption is motivated by the historical alignment of public and private sector wages at the 90th percentile, both in levels and in trends.[11] Institutionally, public sector wages are set based on the recommendations of independent government pay review bodies, which explicitly consider private sector wage trends in their decisions. Prior to 2010, these recommendations were routinely adopted by the government. Between 2010 and 2017, however, they were overridden in favour of the pay cap.[12]

3.1.1 Wage Growth Along the Distribution

To illustrate the impact of the pay caps, Figure 2 plots wage growth along the distribution relative to 2006. The left panel shows the public sector, while the right panel shows the private sector. In both the public and private sectors, wage growth was approximately uniform across the distribution. Wage growth remained positive in the public sector between 2006 and 2010, as the government continued to honour pre-arranged pay agreements following the 2007 recession. In the private sector, however, wage growth remained flat as the wage effects of the financial crisis began to bite in 2010.

The dashed line represents the approximate bite point of the pay caps in each sector. We choose 33 hours per week, as it corresponds to the average number of hours in a standard full-time teacher's contract. This is slightly below the standard 37 hour standard full time contract, but it ensures that we capture an important contingent of public sector workers. In the public sector, this corresponds to the 44th percentile, and in the private sector, the 64th percentile. The difference largely reflects variation in the observable composition of public- versus private-sector workers, which we outline in the next section. Across both the public and private sectors, there is no substantial change in wage growth around the cutoff between 2006 and 2010.

The orange line in Figure 2 represents wage growth between 2006 and 2019, while the

[11] See Table 1 for evidence on pre-2010 public-private equivalence at the 90th percentile.

[12] For example, the 2023 NHS Pay Review Body report (paragraph 3.215) compares the wages of health and social care workers in the public and private sectors. Similar comparisons are featured in each annual report.

difference between the orange and blue lines captures wage growth between 2010 and 2019. In the public sector, wage growth evolved very differently above and below the pay cap threshold. Below the threshold, wages changed relatively little between 2010 and 2019, whereas above the threshold they declined increasingly along the distribution, reaching approximately -4.0% at the 90th percentile. In contrast, although private-sector wage growth was stronger at the lower end of the distribution, there is no clear change in trend around the corresponding cutoff. The figure therefore provides strong suggestive evidence that the pay caps were binding in the public sector, with no comparable pattern around the threshold in the private sector.

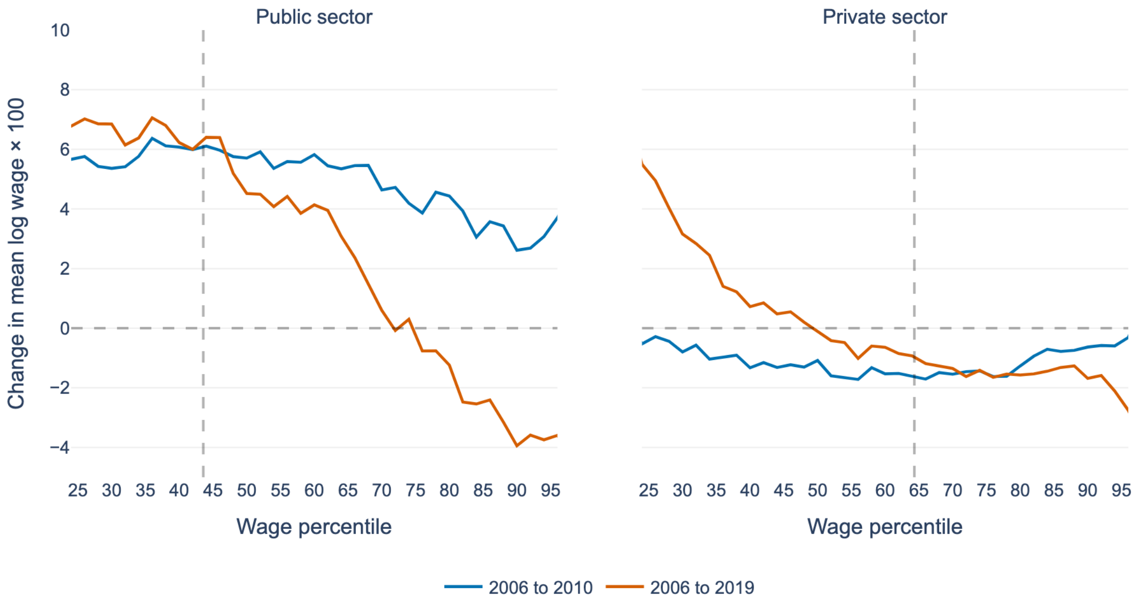


Figure 2: Wage Growth Along the Distribution Relative to 2006

Note: The figure shows real hourly wage growth at equally spaced percentiles, relative to 2006. The vertical dashed line converts the £21,000 annual threshold into an hourly equivalent a 33 hour per week contract. Percentiles below the 25th are excluded to abstract from spillover effects of the living wage; observations above the 95th percentile are omitted. *Source:* ASHE.

3.2 Characteristics of Public Sector Workers

The overall inequality effect of changes to public sector wages and composition depends on where public sector workers are located within the overall wage distribution. This position is partially determined by the observable characteristics of public sector employees. Table 1 presents descriptive statistics for public and private sector workers in 2010 and

2018, respectively.

Table 1: Public and Private Summary Statistics

	Public			Private		
	(1)	(2)	(3)	(4)	(5)	(6)
	2010	2018	% Chg	2010	2018	% Chg
Share of employees	30.8	24.2	-21.3	69.2	75.8	9.5
Degree^{+}	46.9	55.6	18.6	25.7	35.9	39.7
Male	34.8	32.3	-7.2	59.3	58.3	-1.7
Age $\leq$ 25	7.1	7.1	0.0	18.9	17.7	-6.3
(26,55]	79.1	77.2	-2.4	69.7	69.5	-0.3
55+	13.8	15.6	13.0	11.2	12.8	14.3
South of England	55.9	55.5	-0.7	65.0	65.6	0.9
Weekly hours	31.6	31.8	0.6	34.7	34.2	-1.4
Teacher-doctor share	25.8	30.1	16.7	1.3	1.7	30.7
N (ASHE)	47820	36878		107312	120092	
N (LFS)	7909	5998		17075	15825	

Note: The table reports selected summary statistics for individuals employed in the public and private sectors. Hourly wages are deflated to 2019 prices. The South of England is defined using the government office regions London, the South East, the South West, and the East of England. A degree-educated worker is defined as anyone with education beyond A-levels. *Sources*: ASHE and LFS ($^{+}$).

In 2010, relative to the average private sector worker, public sector employees were 21.2pp more likely to hold a degree and were, on average, 3.9 years older. Consistent with prior findings (Postel-Vinay, 2015), these differences suggest that the public sector contains a different composition of human capital relative to the private sector. This reflects, in part, the occupational structure of the UK public sector, where the state is the monopoly provider of healthcare and education—two high-skill professions in which degree-level education is typically required.

3.2.1 Changes over Time

As part of the government's fiscal retrenchment, public sector employers were required to reduce vacancy postings and refrain from renewing temporary contracts (Andrews, 2015; ONS, 2019). Moreover, as the pay caps began to bind, it is natural to consider whether public sector workers may have voluntarily exited the sector in response to the increasing value of outside options. Together, these factors contributed to a 7.0pp decline in the public sector share of total ASHE employment between 2010 and 2019 from 30.8% to 23.8%, as shown in Figure 3.

Columns 3 and 6 of Table 1 summarise how these factors changed the composition of the public sector by reporting the percentage change in observable worker characteristics between 2010 and 2018. Over this period, the average level of education increased in both sectors. However, growth in the share of degree-educated workers was significantly smaller in the public sector than in the private sector. Second, the relative share of men in the public sector declined. Given the large unconditional wage premium that men earn over women, the reduction in the male share likely shifted the public sector wage distribution leftward by reducing the average wage level. Third, there was a disproportionately large reduction in the share of individuals aged 26-55 in the public sector. This cell represents the highest wage age bin, so this reduction likely suppressed wage growth.

In contrast, there was a shift away from part time work in the public sector, with the number of hours increasing by 0.2, while falling by 0.5 in the private sector. This movement away from part time work may have contributed to increased wages in the public sector. Likewise, the occupational composition of the public sector shifted considerably over this period, with employment becoming more concentrated in core public sector occupations, such as teaching and medicine. Owing to the relatively high wages of these professionals, this shift in occupational composition likely worked to increase wages at the top of the distribution. In the results section, we move to understand the exact nature of these compositional changes on wage growth along the distribution.

3.2.2 Public Sector Conditional Premium

Given that public sector workers differ substantially in observable characteristics from their private sector counterparts, comparing raw wage percentiles does not yield a like-for-like comparison. Figure 4a takes the first step towards addressing this by plotting the public sector wage premium—the mean wage differential between public and private sector workers—before and after controlling for observable characteristics.

The unconditional premium is represented by the black line, which is obtained by regressing individual log real wages on year fixed effects and interactions between year fixed effects and the individual's sector (public or private). The coefficients on these inter-

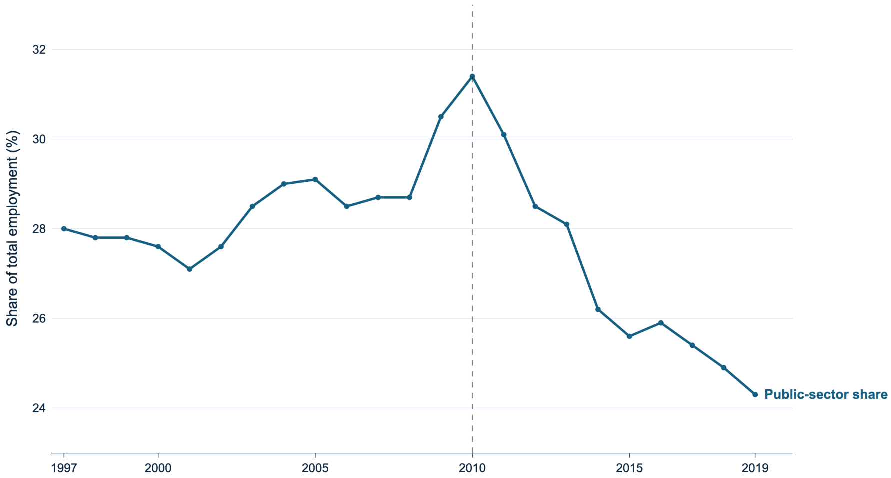


Figure 3: Public Sector Share of Total Employees (1997–2019)

Note: The figure plots the public sector share of total employees between 1997 and 2019. *Source*: ASHE.

actions represent the average public sector wage premium, in log points, relative to the private sector. The average unconditional premium was 20.9 log points, but this varied considerably over time, increasing from 17.7 log points in 2006 to 23.3 log points in 2010 before falling to 19.1 log points in 2019.

The blue line represents the wage premium after conditioning on age squared, sex, and full-time work, alongside individual and region fixed effects (with regions measured at the NUTS 1 level). We estimate:

$$\ln(w_{it}) = \alpha_t \cdot \text{public}_{it} + \beta_t X_{it} + \phi_i + \theta_t + \varepsilon_{it}, \tag{1}$$

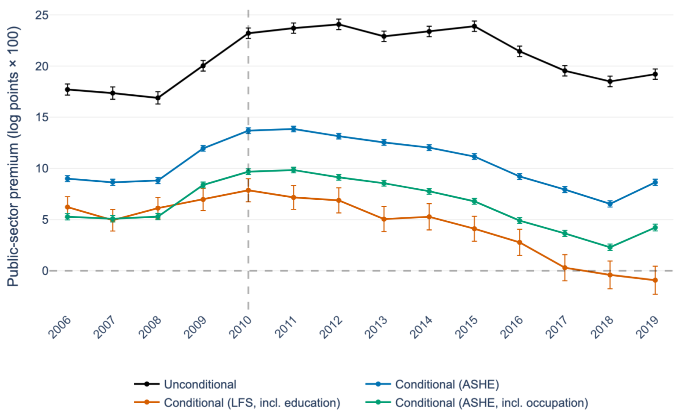
where α_t is the coefficient on an indicator for public sector employment, w_{it} is the hourly wage, X_{it} is the vector of individual characteristics outlined previously, and ϕ_i and θ_t denote individual and time fixed effects, respectively. The coefficients α_t trace the conditional public sector wage premium over time. There was an immediate level shift relative to the unconditional case, with the conditional premium averaging 10.5 log points over our sample. The trend in the conditional premium was similar to that of the unconditional premium, increasing in the lead-up to the caps and decreasing thereafter.

The set of observables included in the conditional premium constitutes our preferred specification throughout the analysis, but our data omit education, a key observable that may drive wage differences between sectors. Although this is largely captured by time-invariant individual fixed effects, we conduct a robustness test using the LFS sample of employees, which allows us to directly calculate the public sector wage premium while controlling for education. Because these data are cross-sectional, we no longer include individual fixed effects. This conditional premium is represented by the orange line and displays a similar broad trend over time, with two notable differences: the average premium was 4.4 log points, indicating a substantial downward level shift, and after 2010 the premium fell at a considerably faster rate, reaching approximately -1 log point in 2019.

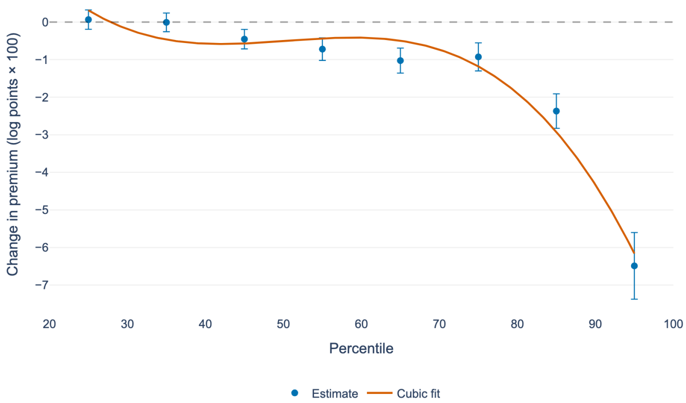
In a second robustness check, we account for the changing occupational composition of the public sector by controlling for two-digit occupation fixed effects.[13] This is not our preferred specification because occupational choice is potentially an outcome on the causal pathway between sectoral choice and wages. However, including these controls allows us to tentatively examine the role of occupational composition in shaping the conditional wage premium. The green line shows this premium. The premium was substantially lower than in the standard conditional specification, indicating that differences in occupational composition accounted for part of the level of the conditional public sector premium. Importantly, however, the occupation-adjusted premium followed an almost identical trend to the standard conditional premium. This suggests that occupational differences explain part of the level of the public sector premium, but that changes in occupational composition contributed little to its evolution over time.

Figure 4b extends the analysis of the public sector premium beyond the mean, considering the evolution of the public sector wage premium at deciles across the residual wage distribution between 2010 and 2019. Our approach involves taking the residuals from equation (1) without including the sectoral regressor. We then group these residuals into deciles in 2010 and 2019 and calculate the change in the premium within each decile. The

[13] Table A1 provides a balance table showing that there is common support between public and private sector workers within each occupation, allowing for identification.



(a) Conditional Premia



(b) Change Along the Distribution (2010–19)

Figure 4: Public Sector Wage Premium Over Time and Across the Distribution

Note: Panel A shows the unconditional public sector wage premium, calculated as the coefficient on a public sector indicator when regressed on log real hourly wages (after accounting for time fixed effects). The conditional premium additionally controls for age (squared) and full-time status, alongside individual and time fixed effects. The LFS premium omits individual fixed effects in favour of education and the final ASHE premium also controls for occupation. Panel B shows changes in the conditional premium at residual wage deciles. The bottom two deciles are omitted to abstract away minimum wage effects.*Source*: ASHE results show that the premium decreased along the wage distribution over our horizon, but disproportionately so at the top decile. This suggests that it was not just specific groups of individuals in the public sector that lost out over this period, rather workers within groups lost out if they were highly remunerated relative to their peers. Table A2 records the premium in each year across each wage decile. Across the bins, we observe marginally faster wage growth at higher wage bins between 2006 and 2010, followed by disproportionately large wage reductions in those higher bins thereafter.

3.2.3 Relative Position of Public Workers in the Distribution

Given the higher average degree share and age of public sector workers relative to their private sector counterparts, among other factors, we might expect public sector employment to be concentrated toward the top of the overall wage distribution. This is indeed the case. In 2010, the public sector accounted for 41.6% of employment at the 75th percentile of the log hourly wage distribution, compared with only 30.1% at the 25th percentile.[14].

By 2018, however, the share of public sector workers had declined along the distribution. This represents both the aggregate shrinkage of the public sector share of total employment and the compression of earnings at the top end. Importantly, the drop was not uniform. Above the pay cap threshold, the public share fell by 8.9pp, compared to just 6.5pp below the threshold. This implies that public sector workers became decreasingly represented at the top of the wage distribution.

3.3 Inflow and Outflow

The decline in the public sector share of employment may reflect either an increase in outflows from the public sector or a reduction in inflows. This subsection examines which of these channels contributed most to the compositional changes described above, drawing on data from both UKLS and ASHE.

[14]The full distribution of the public sector employment share in 2010 and 2018, calculated using 0.1 log-point wage intervals, is shown in Figure A1

Table 2: Transition Matrices

		$t+1$					
		Public	Private	Self Emp.	Unemp.	Retired	Inactive
	Panel A: 2004–2009						
t	Public	88.0	5.0	0.9	1.0	1.9	3.3
	Private	2.6	87.0	2.9	2.4	1.6	3.5
	Self Employment	1.5	11.7	81.4	1.5	1.7	2.1
	Unemployment	6.8	27.5	5.5	32.9	3.5	23.9
	Retired	0.3	0.8	0.2	0.2	95.6	2.9
	Inactive	5.3	12.4	1.5	5.0	5.3	70.5
	Panel B: 2011–2018						
t	Public	79.9	12.1	1.1	1.2	2.9	2.8
	Private	5.8	83.0	3.5	2.4	1.9	3.5
	Self Employment	1.5	12.7	78.7	1.3	3.5	2.3
	Unemployment	4.4	21.9	4.1	42.9	3.5	23.3
	Retired	0.3	0.6	0.5	0.2	96.8	1.6
	Inactive	4.2	12.6	1.7	8.4	4.3	68.9
	Panel C: Difference						
t	Public	-8.1	7.1	0.2	0.2	1.0	-0.5
	Private	3.2	-4.0	0.6	0.0	0.2	0.0
	Self Employment	0.0	1.0	-2.7	-0.2	1.8	0.2
	Unemployment	-2.4	-5.6	-1.4	10.0	0.0	-0.6
	Retired	0.0	-0.2	0.3	0.0	1.2	-1.3
	Inactive	-1.1	0.2	0.2	3.4	-1.0	-1.6

Note: Individuals are classified into six labour market states: public sector employment, private sector employment, self-employment, unemployment, retirement, and inactivity. Panel A reports average annual transition probabilities between states from 2004 to 2009. Panel B reports the same for 2011 to 2018. Each row shows the probability of being in a given state in year $t+1$, conditional on the state occupied in year t; thus, row probabilities sum to 100. Panel C presents the percentage point change in transition probabilities between the two periods. We reweight the sample to produce a balanced cross-section and mitigate composition effects due to ageing. *Source*: UKLS

Table 2 presents a transition matrix of annual labour market flows. Panel A shows the average annual transition probabilities between labour market states from 2004 to 2009. Panel B reports the corresponding probabilities for the post-reform period, from 2011 to 2018. Panel C displays the percentage point change in transition probabilities between the two periods.[15]

Focusing on the public sector in Panel A, 88% of workers remained in public sector employment in the subsequent year during the pre-2010 period. Public sector retention—measured as the proportion of workers remaining in the same labour market state from one year to the next—was higher than for any other labour market state except retirement. Among those who exited the public sector, most transitioned into private sector employment, followed by inactivity. Compared with private sector employees, public sector workers were more likely to remain in employment overall. These patterns underscore the high degree of job security associated with public sector employment prior to 2010. After 2010, outflows from the public sector increased by 8.1pp, while into declined by only 0.3pp. Approximately 85% of the increase in outflows consisted of workers transitioning to the private sector, with most of the remainder moving into retirement. As a result, public sector retention fell to 79.9%, below the private sector retention rate of 83.0%.

These results suggest that post-2010 changes in the composition of the public sector wage distribution were primarily driven by increased outflows rather than reduced inflows. In Figure A2, we decompose inflows and outflows across the wage distribution to assess descriptively whether the pay caps may have shaped selection into or out of the public sector. In each case, we find little evidence of such selection. Inflows remain approximately constant across the distribution when comparing flows in 2009–10 with those in 2017–18. Although outflows increased, the increase was approximately uniform across the distribution.[16]

[15]We omit 2010 from the analysis because it marks the transition from the British Household Panel Survey (BHPS) to Understanding Society (USoC), which introduced a change in the sampling structure. To avoid bias, we also reweight the data to ensure a balanced cross-section and adjust for compositional effects associated with ageing.

[16]Figure A3 repeats this exercise but groups individuals into vigintiles based on their individual fixed effects estimated from equation (1), excluding sectoral indicators. The resulting inflow and outflow patterns closely mirror those obtained using raw wages, providing suggestive evidence that changes in

4 Methodology

4.1 Decomposing Densities

Each worker possesses a vector of observable characteristics, which earn labour market returns through a sector-specific price vector. These characteristics include age, age squared, sex, full time work, and region fixed effects. We denote the characteristics x_t and the returns β_t^s, where $s \in \{\text{public}, \text{private}\}$ and t is time. The returns capture both the relative compression of public sector wages and the variation in returns across sectors, and they vary over time to reflect the wage changes documented in Section 3.1. We denote the sectoral wage distribution as $f_t^s(w)$.

We fix 2010 as the base year and compare it with a terminal year τ, which ranges over the post-reform horizon 2011–19. For each τ, let f_{CF}^s denote the counterfactual density that would prevail if the distribution of characteristics remained at its 2010 level while returns followed their period-τ path. The change in the sectoral wage density between 2010 and τ can then be decomposed as:

$$\underbrace{\Delta f^s(w \mid \beta^s, x)}_{\text{Overall change}} = \underbrace{f_{CF}^s(w \mid \beta_\tau^s, x_{10}) - f_{10}^s(w \mid \beta_{10}^s, x_{10})}_{\text{Wage structure effect } = \Delta f^s(\beta)} + \underbrace{f_\tau^s(w \mid \beta_\tau^s, x_\tau) - f_{CF}^s(w \mid \beta_\tau^s, x_{10})}_{\text{Characteristic effect } = \Delta f^s(x)} \tag{2}$$

The first term is the wage structure effect, capturing the change in the wage density that would have occurred had the composition of the workforce remained at its 2010 level. In the public sector, where pay is largely set by government policy, this reflects institutional changes such as the pay caps and the introduction of the National Living Wage in 2016. It also absorbs any change in the distribution of factors not captured by observable characteristics, and so does not by itself distinguish the pay caps from other shocks to the conditional wage structure. Section 4.3 addresses this by differencing the public sector wage structure effect against its private sector counterpart.

The second term is the characteristic effect, capturing changes in the wage density driven

labour market flows were similar across workers with different estimated individual fixed effects.

by the shifting composition of observable characteristics, holding returns at their period-τ values. This aggregate effect reflects the composition shift owing to hiring freezes, the non-renewal of temporary contracts, and voluntary exits in response to pay restraint. We do not isolate these channels individually.

Isolating the characteristic effect through reweighting requires that the conditional wage density given characteristics is invariant to the change in composition: shifting the distribution of workers across observable types does not itself alter the returns those types command – the standard assumption in the decomposition literature. The decomposition is also sequential, and hence path-dependent: the wage structure effect is evaluated holding composition at its 2010 level, and the characteristic effect holding returns at their period-τ level, we maintain this ordering throughout.

4.2 Calculating Characteristic and Wage Structure Effects

The DiNardo et al. (1996) reweighting approach isolates the characteristic effect by decomposing the wage density at any time t into a conditional wage density given characteristics, $h_t^s(w \mid x)$, which embeds the returns β_t^s, and the distribution of those characteristics, $g_t^s(x)$. The observed period-τ density is:

$$f_\tau^s(w \mid \beta_\tau^s, x_\tau) = \int h_\tau^s(w \mid \beta_\tau^s, x)\, g_\tau^s(x)\, dx \tag{3}$$

The counterfactual density $f_{CF}^s(w \mid \beta_\tau^s, x_{10})$ from Equation (2)—the density that would have prevailed in period τ had characteristics remained at their 2010 level while returns followed their period-τ path—is obtained by reweighting the observed period-τ density so that its characteristic distribution matches that of 2010:

$$f_{CF}^s(w \mid \beta_\tau^s, x_{10}) = \int h_\tau^s(w \mid \beta_\tau^s, x)\, g_\tau^s(x)\, \phi^s(x)\, dx \tag{4}$$

The reweighting function $\phi^s(x)$ is the ratio of the 2010 to the period-τ characteristic density. Applying Bayes' rule expresses it through the probability of period membership conditional on characteristics:

$$\phi^s(x) = \frac{g^s_{10}(x)}{g^s_\tau(x)} = \frac{\Pr(t=\tau)}{\Pr(t=10)} \cdot \frac{\Pr(t=10 \mid x)}{\Pr(t=\tau \mid x)} \tag{5}$$

Characteristics more common in 2010 than in period τ receive a weight above one, so that the reweighted period-τ distribution of characteristics matches that of 2010. We estimate the conditional period probabilities with a logit on observable characteristics, and recover the three densities—observed 2010, observed period-τ, and counterfactual—non-parametrically using kernel methods. The reweighting requires common support: characteristic combinations present in 2010 must also appear in period τ. Applying the identity in Equation (2) then yields the wage structure and characteristic effects.

To identify which observables account for the aggregate characteristic effect, we further decompose the reweighting sequentially across covariates following DiNardo et al. (1996). Starting from the period-τ distribution, we first reweight one characteristic to match its 2010 distribution, then reweight a second characteristic conditional on the first, and continue until the joint distribution of the full set of observables matches that of 2010. Each step therefore constructs an intermediate counterfactual that differs from the previous one only through the additional characteristic being reweighted. The contribution of characteristic j is the change in the relevant distributional statistic between these adjacent counterfactuals, so that the sequential contributions sum to the total characteristic effect. Throughout, we maintain the ordering age and age squared, sex, full-time status, then region.

4.3 Capturing Relative Price Changes

The wage structure effect in Equation (2) compares public sector returns in period τ with their own 2010 level. This is not the effect of the pay caps, because public returns would likely not have remained at their 2010 level in the absence of the policy. Private sector returns fell markedly between 2010 and 2014 following the financial crisis, and public and private wages moved together prior to 2010 (Section 3.1). We therefore take the path of private sector returns as the counterfactual for public returns absent austerity. Because returns did not shift uniformly across worker types, the counterfactual must apply the private-sector change in conditional returns to the *public* workforce. We reweight the private period-τ and 2010 densities so that each matches the public 2010 covariate distribution g_{10}^{pub}, and define the relative price effect as the resulting private wage-structure change net of its public counterpart:

$$\Delta f_{rel}^{pub}(\beta) = \underbrace{\int \left[h_{\tau}^{priv}(w \mid \beta_{\tau}^{priv}, x) - h_{10}^{priv}(w \mid \beta_{10}^{priv}, x)\right] g_{10}^{pub}(x)\, dx}_{\Delta f_{R}^{priv}(\beta)} - \Delta f^{pub}(\beta). \qquad (6)$$

Standardising the private sector to the same 2010 population as the public sector ensures the contrast reflects the difference in returns rather than the difference in workforce composition across sectors.[17]

Each term is estimated by the reweighting procedure of Section 4.2.[18] The construction shares the intuition of a conditional difference-in-differences design, comparing changes in the public-sector wage structure with those in the private sector over the same period. Identification requires a distributional parallel-paths assumption: absent public-specific wage-setting interventions, the conditional public-sector wage density would have evolved along the same path as the private-sector density. Under this assumption, common structural shifts in the labour market—such as the 2016 National Living Wage and broad shifts in labour demand—are accounted for, provided they reshape the wage distribution similarly for comparable workers across both sectors. The relative price effect should consequently be interpreted as the deterioration in public-sector returns relative to the private-sector counterfactual. We then use the institutional design of the policy and additional threshold-based evidence to assess whether the pay caps were the principal driver of this divergence

[17]When the public and private 2010 covariate distributions coincide, so that $g_{10}^{pub} = g_{10}^{priv}$ and the weights are one, the first term reduces to $\Delta f^{priv}(\beta)$ and Equation (6) collapses to the naive difference of sector wage-structure effects, $\Delta f^{priv}(\beta) - \Delta f^{pub}(\beta)$. The composition adjustment therefore binds only to the extent that the two workforces differ in observables.

[18]In all cases our counterfactuals produce monotonic valid CDFs over the corresponding wage support.

The most direct evidence for our assumption is that the post-2010 divergence between public and private wage growth is concentrated above the £21,000 threshold at which the caps bind, rather than below it (Figures 1 and 5). Below-threshold public workers are exposed to the same economy-wide shocks as those above but not to the binding cap, and serve as an effective within-sector placebo: their wage growth tracks the private sector throughout, whereas the divergence emerges only where the policy bites. We show this formally in a triple difference event study framework in Appendix A, where we compare the relative wage trajectories of marginally capped to uncapped workers in the public sector relative to the private sector over time. We show clear evidence that wages declined in the public sector above the cutoff without evidence of a pre-trend, and we find limited evidence of any private sector response until at least 2015. Our results are resoundingly robust when we consider workers further above the pay cap threshold, providing strong evidence that the caps were responsible for the wage declines in the public sector.

4.4 Mapping to the Overall Wage Distribution

The overall wage distribution is a mixture of the public and private sector distributions. Let δ_t denote the public sector share of employment at time t; the overall density is:

$$f_t^{all}(w \mid \beta_t^{all}, x_t) = \delta_t \, f_t^{pub}(w \mid \beta_t^{pub}, x_t) + [1 - \delta_t] \, f_t^{priv}(w \mid \beta_t^{priv}, x_t) \tag{7}$$

The two sector densities differ along the wage distribution: the public share of employment peaks near the 75th percentile and is thin at the bottom (Figure A1). Reallocating employment between the sectors therefore reshapes the overall distribution unevenly, concentrated where the two densities are furthest apart.

Differencing the mixture between 2010 and τ yields three sources of change: movements in the public density, movements in the private density, and movements in the employment share. Combining these with the within-sector decompositions of Sections 4.1–4.3 gives:

$$\begin{aligned}\Delta f^{all}(w) = {} & \delta_{10} \underbrace{\left[\Delta f_R^{priv}(\beta) - \Delta f_{rel}^{pub}(\beta)\right]}_{\text{Public wage structure}} \\ & + \delta_{10} \quad \underbrace{\Delta f^{pub}(x)}_{\text{Public characteristics}} \\ & + [1-\delta_{10}] \underbrace{\Delta f^{priv}(w)}_{\text{Private shifts}} \\ & + \underbrace{(\delta_\tau - \delta_{10}) \left[f_\tau^{pub}(w) - f_\tau^{priv}(w)\right]}_{\text{Share effect}}\end{aligned} \tag{8}$$

Each within-sector change is evaluated holding the employment share at its 2010 level, δ_{10}, and the share effect holding the sector densities at their period-τ values.

The public wage structure term separates the private returns path from the relative price effect of Equation (6); the public characteristic term captures shifts in the distribution of observable characteristics within the public sector.

To recover the conditional inequalities of Section 5, we repeat the decomposition separately for men and women, and for the South and the rest of the UK, estimating sector-specific densities within each subgroup.

Equation (8) treats private sector wages and composition as unresponsive to public pay restraint—a partial equilibrium assumption. Three considerations support it. First, the public labour force is roughly one-third the size of the private sector, so even a large outflow of public workers raises private labour supply only modestly. Second, compositional change contributes little to private sector wage growth over the period (Section 5, Figure 6). Third, Bradley et al. (2017) find that private wage offers respond only marginally to public pay cuts: a cut reducing the public wage bill by 10%—comparable to the fall in real public wages over the period—moves private offers by less than 1%.

4.5 Estimating the Public Sector Share Effect

The share term in Equation (8) provides a useful accounting benchmark, but it treats the change in public employment as proportional across the observed workforce. In practice, the probability of public sector employment may have changed differently across regions and demographic groups. We therefore supplement the aggregate share term with a richer DiNardo et al. (1996) estimate of the sector-allocation effect.

Let d be a binary indicator equal to one if the worker is employed in the public sector, and let x denote the complete set of observable characteristics. The public-share weight restores the probability of public sector employment conditional on x to its 2010 level, while the distribution of characteristics and returns follow their observed period-τ values. The reweighting function is:

$$\theta(d \mid x) = d \cdot \frac{\Pr(d = 1 \mid x, t = 10)}{\Pr(d = 1 \mid x, t = \tau)} + (1 - d) \cdot \frac{\Pr(d = 0 \mid x, t = 10)}{\Pr(d = 0 \mid x, t = \tau)} \tag{9}$$

Applying $\theta(d \mid x)$ to the period-τ workforce yields a counterfactual distribution in which observationally comparable workers are allocated between the public and private sectors according to the conditional sector probabilities prevailing in 2010. Let $g_\tau(d, x)$ denote the joint distribution of sector and characteristics and $h_\tau(w \mid d, x)$ the conditional wage density. The counterfactual density is:

$$f^{all}_{CF,share}(w) = \sum_{d \in \{0,1\}} \int h_\tau(w \mid d, x)\, g_\tau(d, x)\, \theta(d \mid x)\, dx \tag{10}$$

The public sector share effect is then:

$$\Delta f^{all}_{share}(w) = f^{all}_{\tau}(w) - f^{all}_{CF,share}(w) \tag{11}$$

Unlike the aggregate share term in Equation (8), this measure allows changes in public employment to differ across worker types. It therefore captures changes in sector allocation correlated with region and demographic characteristics, including the disproportionate exit of prime working age public workers documented in Section 3.2.

We estimate the conditional sector probabilities with a logit on the complete set of observables and apply the resulting weights to the period-τ wage distribution. As with the characteristic reweighting in Section 4.2, the procedure requires common support: worker types with positive probability of a given sector assignment in 2010 must also have positive probability of that sector assignment in period τ.[19]

4.6 Inference

We construct pointwise confidence intervals for the counterfactual estimates using a bootstrap procedure following Chernozhukov et al. (2013) with 1000 replications. In each bootstrap replication, observations are resampled and the complete counterfactual procedure is re-estimated, including the reweighting function and resulting counterfactual wage distribution. We then recalculate the relevant counterfactual quantiles and percentile differentials. The empirical distribution of these bootstrap estimates is used to construct pointwise standard errors and confidence intervals separately for each reported percentile and year. Chernozhukov et al. (2013) establish the asymptotic validity of bootstrap inference for counterfactual distribution and quantile functionals under standard regularity conditions.

5 Results

Next, we examine the impact of the austerity measures on inequality between 2010 and 2018 by combining the counterfactual methods outlined above. We begin by assessing the effects of changes in prices and composition across the public wage distribution, before assessing their impact on the overall P90–P50 ratio alongside the commensurate decline in the public sector share. After demonstrating the progressive effects of the policy along the

[19] Although this procedure formally restores the 2010 conditional probabilities of public-sector employment, it yields a counterfactual aggregate share nearly identical to the actual 2010 aggregate share (see Figure A4).

distribution, we consider their implications for conditional inequalities. In particular, we show that the relative price effect spilled over onto the gender pay gap and substantially widened the South's wage premium over the rest of the UK.

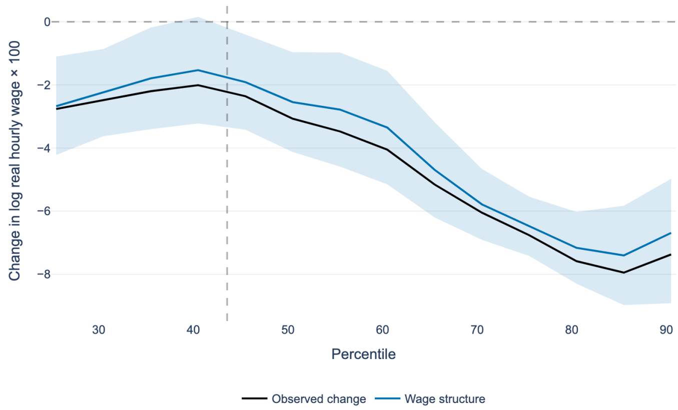
5.1 Relative Prices, Characteristics and the Public Share

Figure 5a illustrates the effects of the pay caps on the public sector by plotting observed and counterfactual real wage growth along the distribution between the 25th and 90th percentiles over the policy period.[20] The wage structure effect led to a reduction in wages along the distribution, which was approximately constant at -2.0% until the cap threshold. Above the threshold, wages began to decline at an increasing rate, reaching a trough of -7.6% in the top 85th percentile of the wage distribution. The accompanying confidence intervals, constructed following Chernozhukov et al. (2013), confirm that the wage structure effect was significantly different from zero and increased in absolute magnitude along the distribution. The wage structure effect closely mirrored observed wage growth over this period, indicating that changes in the wage structure were the dominant factor driving wage growth in the public sector.

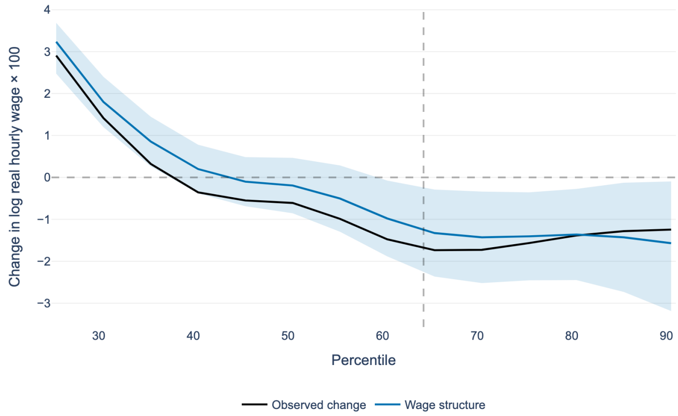
Contrasting the public sector with the private sector in Figure 5b, wage growth was again declining along the distribution, although to a much lesser extent than in the public sector. At the cap threshold, the private sector wage structure effect reduced wages by 1.2%, falling to 1.5% at the 90th percentile, although this difference is not statistically significant. As in the public sector, observed wage growth closely mirrors the wage structure effect, confirming that changes in the wage structure dominated wage growth in both sectors over this period.

The relative price effect compares the wage structure effect in the private sector with that in the public sector, conditional on observable characteristics, and then integrates the difference over the distribution of characteristics in the public sector base year, 2010, to generate the unconditional wage counterfactual. Figure 5c plots the relative price

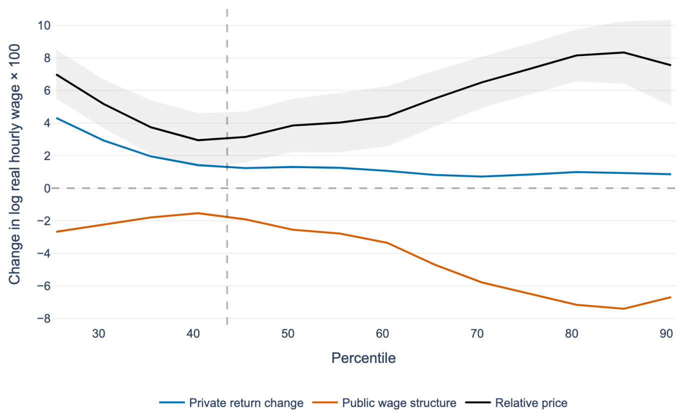
[20]We drop observations below the 25th percentile to abstract away from minimum wage effects.



(a) Public



(b) Private



(c) Relative Price Effect

Figure 5: Wage Structure and Relative Price Effects (2010–18)

Note: The figure presents wage structure and relative price effects in the public and private sectors, estimated using a DiNardo et al. (1996) decomposition that controls for sex, age (squared), full-time status, and region fixed effects. Panel A presents the public sector wage structure effect alongside observed wage growth, with 95% confidence intervals estimated following Chernozhukov et al. (2013). Panel B presents the corresponding results for the private sector. Panel C presents the relative price effect in black, with standard errors, alongside its components: the private sector wage structure effect integrated over the distribution of public sector characteristics in 2010, net of the public sector wage structure effect. *Source:* ASHE.

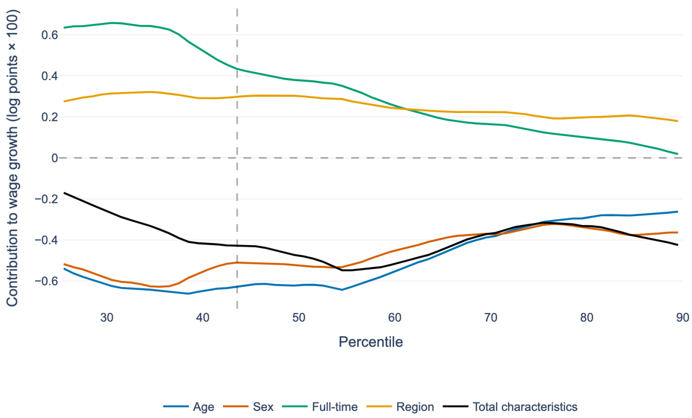
effect alongside its constituent components along the public sector wage distribution. The relative price effect was positive and increased significantly along the distribution: at the 90th percentile, public sector wages would have been 7.6pp higher, compared with 3.1pp at the bite point. The private sector component was approximately zero at all percentiles above the cap threshold, suggesting that nearly all of the relative price effect is driven by declines in the returns to observable characteristics in the public sector. The concentration of the relative price effect above the policy threshold, together with the negligible private-sector component at these percentiles, is consistent with the pay caps being the principal driver of the post-2010 deterioration in public-sector relative returns.

Our data do not include education by default. Consequently, time-varying changes in the returns to, or composition of, education may be absorbed by the estimated wage structure effect, potentially contaminating our estimates given the shifts in educational composition within each sector outlined in Section 3.2. We take three steps to address this concern and show that our results are robust. Each is represented as a relative price effect in Figure A5. The first and most direct approach is to replicate our counterfactual using LFS data with a categorical measure of education comprising degree or equivalent, tertiary education below a formal degree, A-level, GCSE, other qualification, and no qualification.[21] The second approach proxies for education in ASHE using individual fixed effects, which we calculate over the pre-policy period to avoid contamination from the pay cap policy. We estimate these individual fixed effects using log hourly wages as the outcome and our standard set of observable characteristics, excluding sector. While this approach necessarily conditions on workers observed in the labour market during the pre-policy period, the resulting relative price effect closely matches the trend observed in both the baseline specification and the LFS.

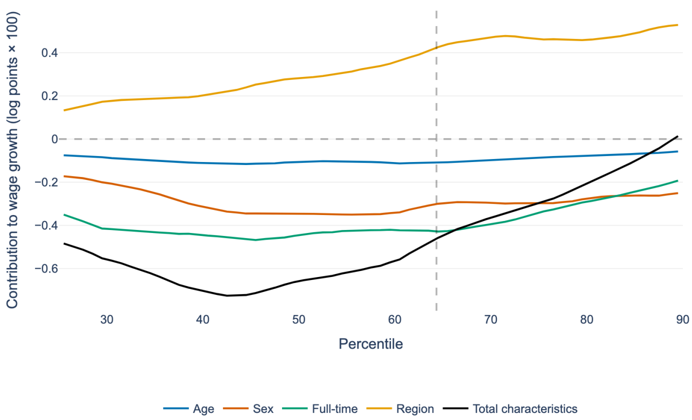
The final robustness check leverages the longitudinal nature of our data to proxy for education using age-of-entry bins. We define two indicators: one for workers who first enter the data between ages 21 and 25, and another for those who enter between ages 16

[21] We prefer ASHE for our main specification owing to its significantly larger sample size and lower propensity for self-reporting bias and selective non-response.

and 20. This approach is motivated by the ages at which graduates and non-graduates typically enter the labour market. We deliberately leave workers who first enter the sample after age 25 outside either category because our data only extend back to 1997, meaning that their observed age of entry is unlikely to correspond to their true labour-market entry age. The resulting relative price effect is likewise approximately flat below the cap threshold before increasing substantially along the distribution.



(a) Public



(b) Private

Figure 6: Public and Private Sector Characteristic Effects (2010-18)

Note: The figure presents characteristic effects in the public and private sectors, estimated using a DiNardo et al. (1996) decomposition that controls for sex, age (squared), full-time status, and region fixed effects. Each panel includes an overall estimate in black as well as the contribution of each characteristic. *Source:* ASHE.

The difference between the wage structure effect and observed wage growth comes from the role of characteristics. Figure 6 depicts the characteristic effects for the public and private sectors on wage growth along their respective distributions between 2010 and 2018. Alongside the overall characteristic effect, we report the individual contribution of each covariate following DiNardo et al. (1996) and we omit standard errors for clarity. In terms of the overall characteristic effect, the changes are broadly similar in both the public and private sectors, except in the top quartile of the distribution, where characteristic changes continued to suppress public sector wages while having no effect on the private sector. Unpacking the composition effect tells a very different story: relative to the private sector, changes to both the age and gender distributions significantly reduced wage growth along the distribution. The results are consistent with our observational evidence that the public sector saw a reduction in male and prime-age workers over the policy period. In contrast, the increase in full-time work in the public sector increased wage growth, in stark contrast to the private sector, where increasing part-time work and zero-hours contracts put downward pressure on the private sector wage distribution. In both sectors, shifts towards employment in high-wage regions increased relative wage growth. This effect is particularly strong at the top of the private sector wage distribution, driving the differential effect observed in the top quartile of the distribution. Even after accounting for the differential effects across observable characteristics, our results point to the wage structure effect as the dominant force shaping within-sector wage inequality over our sample.[22]

Given that public sector workers earned a positive conditional wage premium over their observationally equivalent private sector counterparts throughout our sample, the aggregate reduction in the public sector share of employees is likely to have placed additional downward pressure on the overall wage distribution. Figure 7 examines this by reweighting the public sector share to its 2010 level, while allowing the wage structure and all other covariates to follow their observed paths. The reduction in the public sector share led to a significant and persistent negative effect of approximately 1.8% along the first

[22]We omit the teacher and doctor shares from our preferred characteristics counterfactual because these occupations are relatively rare in the private sector, making the corresponding relative price counterfactual estimates unstable. As a robustness exercise, Figure A6 replicates the characteristics counterfactual while accounting for the shares of teachers and doctors. Their increased concentration contributed 2.1% to wage growth in the public sector, compared with 0.2% in the private sector. However, this contribution was approximately uniform across the wage distribution in both sectors, implying little effect on inequality.

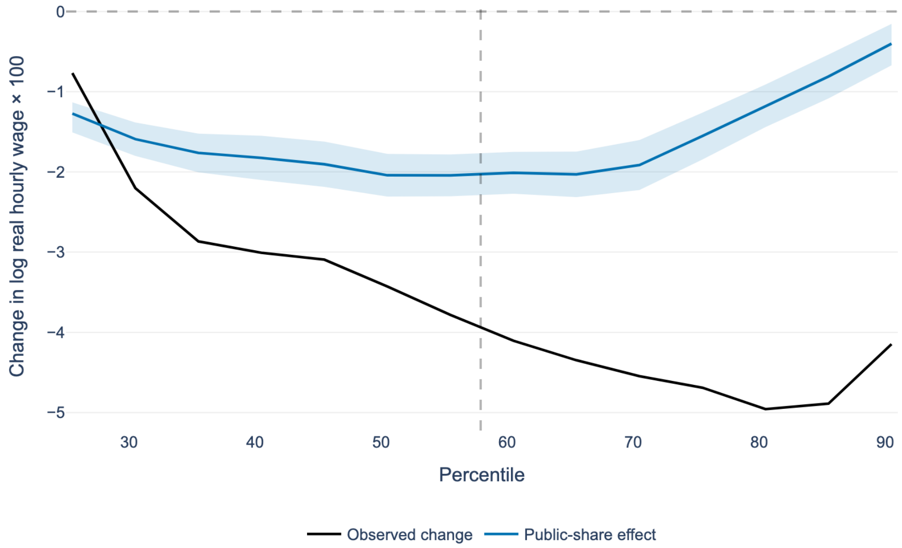


Figure 7: Overall Share Effect (2010–2018)

Note: The figure shows observed and counterfactual wage growth between 2010 and 2018. The black line represents observed wage growth; the blue line represents contribution of the change in the public sector share to that wage growth. The counterfactual is calculated via a DiNardo et al. (1996) decomposition, which conditions on sex, age (squared), full time status, and region fixed effects. We report 95% confidence intervals following Chernozhukov et al. (2013). *Source*: ASHE.

two terciles of the distribution. However, the effect diminishes towards the top of the distribution because the public sector wage distribution is compressed relative to the private sector. As a result, the reduction in the public sector share lowered wages more strongly around the middle of the distribution than at the top, thereby increasing upper-tail wage inequality and offsetting part of the inequality-reducing effect of the pay caps.

To examine whether the pay caps were associated with selective changes in inflows and outflows along the wage distribution, we conduct a robustness exercise that treats inflow and outflow as binary observable characteristics. The inflow indicator equals one if a worker entered the sector in that period, while the outflow indicator equals one if the worker leaves the sector in the following period. In both the public and private sectors, changes in inflow patterns contributed negatively to wage growth between 2010 and 2018.[23] Changes in outflow patterns contributed modestly positively in the public sector and negatively in the private sector. In both sectors, however, the contribution of worker

[23]We end this analysis in 2018 because outflow is necessarily defined using sectoral status in the following period.

flows was small compared to the relative price effect. Moreover, in the public sector, there is little variation in the contribution of either inflows or outflows around the percentile at which the pay caps begin to bind. Figure A7 presents the full decomposition.

5.2 Overall Inequality

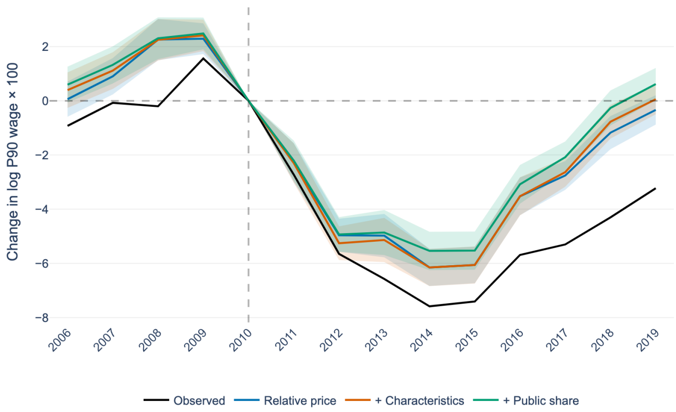
We map each of our effects onto the overall wage distribution following Equations (8) and (9). While the previous figures have focused on counterfactual wage growth between 2010 and 2018, Figure 8 derives observed and counterfactual wages at the 50th and 90th percentiles of the overall distribution for each year beginning in 2006. The purpose of considering the pre-period is to verify that the pay caps generated a progressive pattern of public sector pay adjustment that was not present prior to the policy, thereby providing evidence that our relative price effect captures the impact of the pay caps. We focus on the P90-P50 ratio in order to avoid the contemporaneous impacts of the living wage. The overall P50 has the additional benefit of being just below the bite point of the pay caps in the overall distribution.

Focussing first on the relative price effect, across both the P50 and the P90, wages would have been higher in the post-period under the counterfactual scenario. By 2019, the P50 would have been 0.6pp higher, compared with 2.9pp at the P90. The larger effect at the P90 reflects the progressive incidence of the pay caps and suggests that they played a significant role in shaping not just within-sector inequality but also overall wage inequality. Naturally, pre-period wages would also have been higher under the relative price counterfactual because 2010 represented a high point in public sector pay relative to the private sector. The key distinction is that the counterfactual wage increase would have been approximately equal across the P90 and P50 during the pre-period. This can be seen by turning to the P90-P50 ratio. Prior to the pay cap policy, the relative price effect had a negligible effect on overall inequality. However, in the post-period, counterfactual wage inequality remained approximately static, while observed inequality fell by 1.9pp.

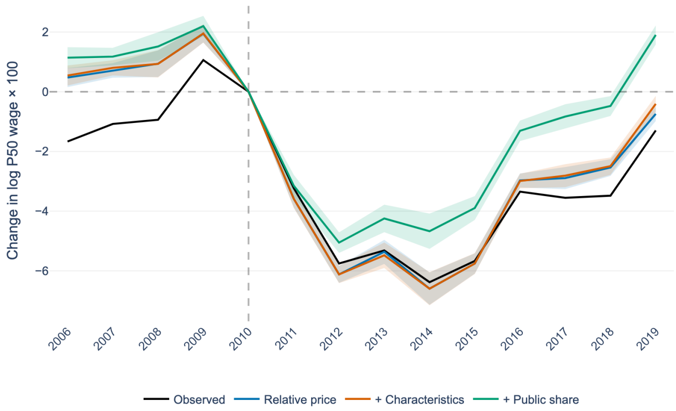
The relative price effect dominates the characteristic effect in every time period and at both percentiles. Reading the figure sequentially, the difference between the relative price counterfactual and the characteristic counterfactual gives the magnitude of the characteristic effect. By the final year of our analysis, the characteristic effect raises counterfactual wages by 0.1pp at the P50 and P90. The similarity of these effects across the two percentiles means that changes in composition played only a limited role in the evolution of overall inequality.

Turning to the public share effect, our results exactly mirror Figure 7, with the counterfactual overall median rising significantly while there is little net effect at the P90. The net effect on overall inequality was that the contemporaneous reduction in the public sector share worked to offset the reduction in inequality generated by the relative price effect. In magnitude, the decline in the public sector share offset approximately 70% of the inequality reduction generated by the relative price effect by 2019.

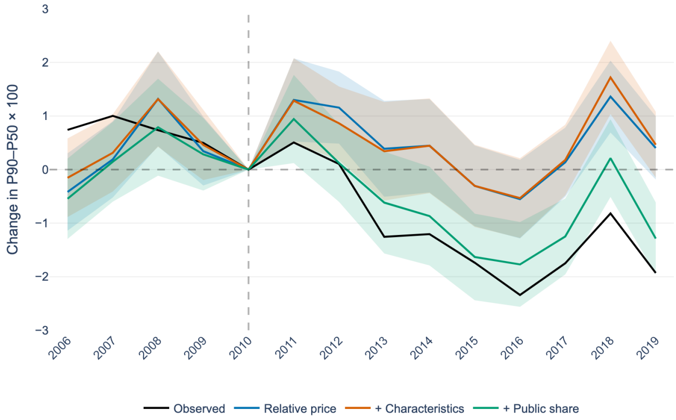
We subject our results to a range of robustness checks, both in terms of outcomes and timing. To address the concern that our analysis is conducted using hourly wages when the caps were implemented using annual salaries, we replicate our analysis using annual wages and full-time-equivalised annual wages. We construct the latter by rescaling annual wages for workers working part time by 33 divided by their observed weekly hours. Additionally, we replicate our analysis using hourly wages net of overtime to address concerns that workers may have supplemented their capped income through additional overtime. In terms of timing, we recognise that 2010 represented a high point for public sector pay relative to the private sector. To address this, we replicate our analysis using the average of 2008–2010 as the base period. Across all robustness checks, our core results hold: the relative price effect drove down inequality, while the reduction in the public sector share partially offset this effect, and composition played a limited role. The results of these checks can be found in Figure A8.



(a) P90



(b) P50



(c) P90-P50

Figure 8: Counterfactual P90–P50 Inequality for the Whole Economy

Note: The figure shows observed and counterfactual wage inequality, measured by the P90–P50 differential. Differences between successive lines capture the relative price effect (black–blue), changes in public sector characteristics (blue–orange), and changes in the public sector employment share (orange–green). Counterfactuals follow DiNardo et al. (1996), conditioning on sex, age (squared), full-time status, and region fixed effects. We report 95% confidence intervals following Chernozhukov et al. (2013). *Source*: ASHE.

5.3 The Gender Pay Gap

Our next exercise traces out the differential magnitudes of our counterfactuals across men and women, allowing us to examine changes in the unconditional gender pay gap. The analysis is motivated by the observation that 65% of public sector employees were women in 2018, compared with only 48% of private sector employees. This suggests that average female wages are disproportionately exposed to shifts in public sector wage-setting practices relative to those of their male counterparts. Similarly, changes in the public sector share are likely to have disproportionate effects on female wages because women are substantially over-represented in public sector employment.

As with the overall wage percentiles, the relative price effect played a much larger role in shaping wages over the sample than the characteristic effect. Between 2010 and 2019, female (male) wages would have been 1.6pp (0.7pp) higher in the absence of relative price shifts and 0.2pp (0.1pp) higher in the absence of characteristic changes.[24] This pattern is not confined to the post-policy period. In the pre-policy period, wages would also have been significantly higher for both men and women in the absence of relative price shifts, reflecting the fact that 2010 represented a relative high point for public sector pay. Importantly, because women were disproportionately concentrated in the public sector, the improvement in relative public sector pay in the build-up to 2010 disproportionately benefited female wages. The opposite occurred following 2010, when the deterioration in relative public sector pay disproportionately suppressed female wage growth.

Taken together, our results suggest that public sector wage setting and the role of the public sector as an employer acted to mitigate gender inequalities in the build-up to 2010 but exacerbated these inequalities in the post-policy period. Figure 9 reinforces this result by plotting the unconditional gender pay gap under our observed and counterfactual scenarios. Focusing on the post-policy period, we find that, in the absence of relative price and composition shifts, and with the public sector share held constant, the gender pay gap would have been 1.9pp lower. Unlike their effects on overall P90-P50 inequality, where the

[24]See Figure A9 for the counterfactuals derived separately for both men and women.

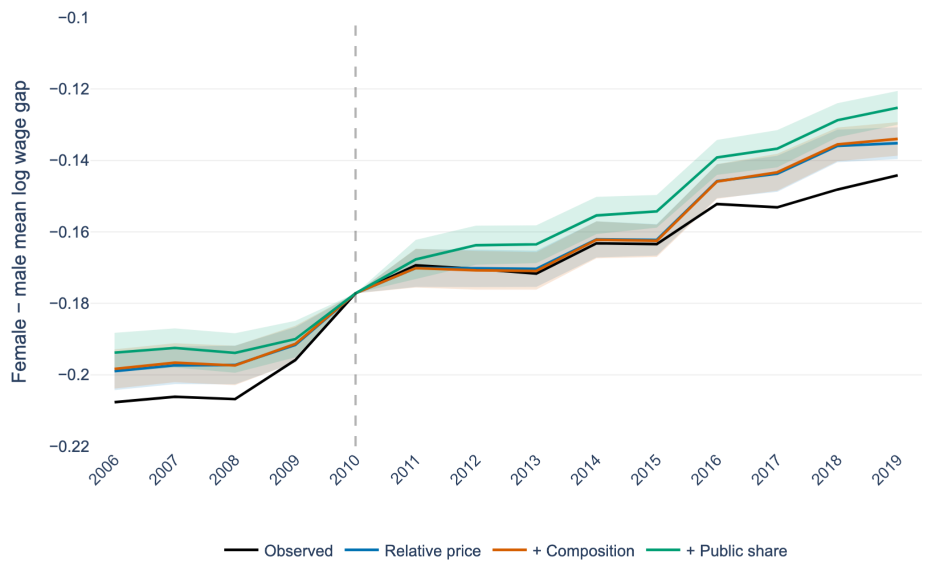


Figure 9: Counterfactual Gender Pay Gap

Note: The figure shows observed and counterfactual gender wage inequality, measured as the mean wage gap between men and women. Differences between successive lines capture the relative price effect (black–blue), changes in public sector characteristics (blue–orange), and changes in the public sector employment share (orange–green). Counterfactuals follow DiNardo et al. (1996), conditioning on sex, age (squared), full-time status, and region fixed effects. We report 95% confidence intervals following Chernozhukov et al. (2013). *Source*: ASHE.

relative price effect reduced inequality, the relative price, characteristic, and share effects all acted to widen gender inequality over this horizon, contributing 47%, 6%, and 47% to the overall widening of the gender pay gap, respectively. The distributional consequences of changes in public sector wage setting therefore depend critically on the dimension of inequality considered: while the pay caps compressed the upper half of the overall wage distribution, they widened gender wage inequality because women were disproportionately exposed to the deterioration in relative public sector pay.

5.4 Regional Inequality

Our final counterfactual analysis explores the regional effects of the pay caps through the lens of the South's wage premium over the rest of the UK. Throughout this section, we define the South as London, the South East, the South West, and the East of England, and compare it with the rest of the UK. Regional inequality in the UK is substantial: as noted by McCann (2020), the UK is one of the most regionally unequal industrialised countries in terms of GDP per capita, productivity, and disposable income. Closing regional wage gaps has been a longstanding policy priority in the UK and, more recently, has been framed under the "Levelling Up" agenda. Initiatives to reduce regional disparities have included devolving powers to the Scottish and Welsh governments and relocating government departments and public bodies to areas outside the South.[25] Reducing regional inequality is also important from a welfare perspective, as improving access to high-quality jobs can increase aggregate output through higher productivity (Bell et al., 2023; Stansbury et al., 2023).

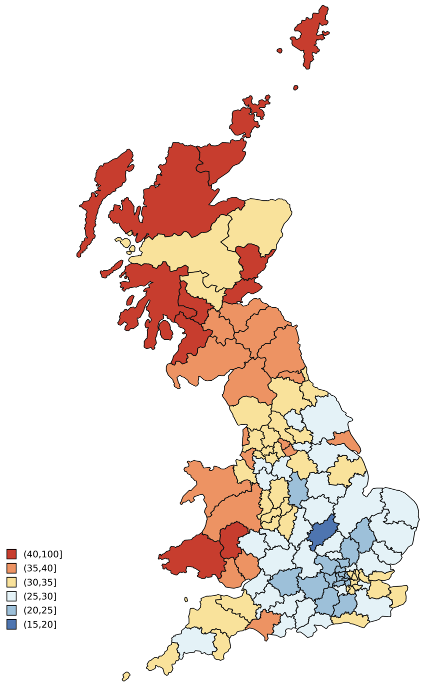


Figure 10: Public Sector Share by Region (2010)

Note: Based on ASHE sample of public sector workers aged 16 and over. The figure shows the public sector share of employment at the postcode region level. The public share refers to the percentage of employed workers in the public sector in 2010.

The government's focus on reducing regional disparities, alongside the disproportionately large volume of private sector investment in the South, has contributed to an elevated share of public sector employment outside the South. Figure 10 illustrates this by plotting the share of public employment by postcode region. The public sector share was highest in Scotland and Wales, peaking at 43% in Swansea. Even in Central London—the seat

[25]For example, the Office for National Statistics relocated its headquarters to Newport, Wales.

of the UK government—the public sector share was just 28%, below the 31% national average. Workers outside the South were therefore disproportionately exposed to changes in public sector wage setting and employment over our sample.

Given this disproportionate exposure, we hypothesise that changes in the public sector wage structure and employment share affected the regional wage gap in much the same way as the gender pay gap. Between 2010 and 2019, wages outside the South would have been 3.6pp higher in the absence of the wage structure, characteristic, and share effects.[26] The same is true, but to a lesser extent, in the South, where wages would have been 2.9pp higher under our counterfactual scenario. Again, the relative price and share counterfactuals dominate, with characteristics playing a muted role.

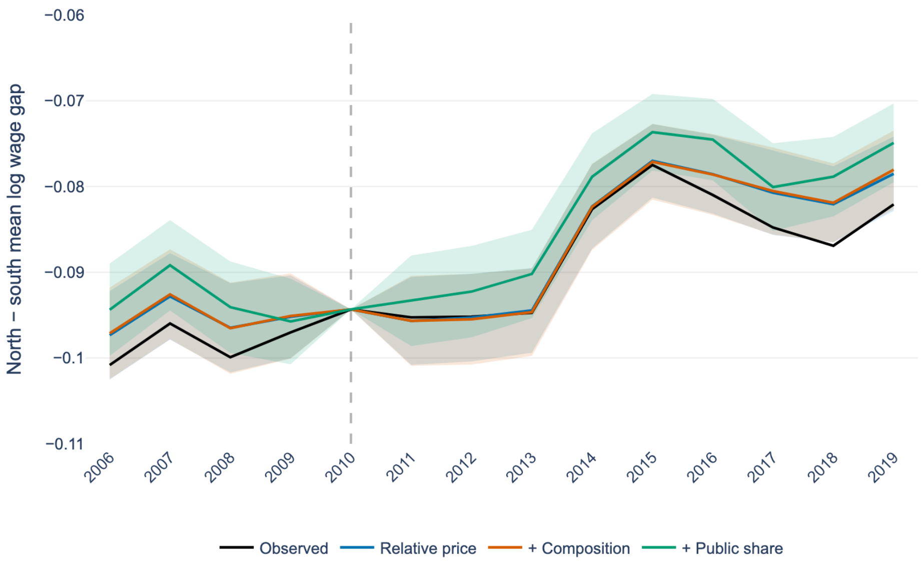


Figure 11: Counterfactual Regional Wage Gap

Note: The figure shows observed and counterfactual regional wage inequality, measured as the mean wage gap between the South and the rest of the UK. Differences between successive lines capture the relative price effect (black–blue), changes in public sector characteristics (blue–orange), and changes in the public sector employment share (orange–green). Counterfactuals follow DiNardo et al. (1996), conditioning on sex, age (squared), full-time status, and region fixed effects. We report 95% confidence intervals following Chernozhukov et al. (2013). The South comprises London, the South East, South West, and East of England. *Source*: ASHE.

Figure 11 maps our results to the regional wage gap. After accounting for all counterfactual effects, the regional wage gap would have been 0.7pp smaller than observed, with the

[26]Figure A10 depicts the counterfactuals for each region.

relative price effect accounting for 50% of the variation, characteristics for 6.5% and the public share for 44%. While the relative price effect alone is not significant at the 95th percentile in 2019, it is significant at the 90th, and the combined effect of all three channels is clearly significant. Again in contrast to the P90-P50 results, the relative price and share effects both worked to widen regional inequality and were approximately equal in magnitude. The distributional consequences of public sector wage restraint therefore depend critically on the dimension of inequality considered: while the pay caps compressed the upper half of the overall wage distribution, they widened both gender and regional wage inequalities because women and workers outside the South were disproportionately exposed to the deterioration in public sector relative pay.

6 Conclusion

This paper studies the distributional consequences of public sector pay caps, a common policy across developed countries in the wake of the global financial crisis. We focus on the particularly binding case of the UK, where all public employees earning above £21,000 in 2010 — approximately 15% of total employees — had their annual wage growth capped to a maximum of 1% until 2017. We construct a counterfactual public sector wage distribution whereby the returns to public sector characteristics follow the private sector over the policy period, which we combine with the observed private sector distribution to study overall inequality. Between 2010 and 2019, top-tail wage inequality in the UK, as proxied by the P90–P50 differential, decreased by 1.9pp, reversing a long-run upward trend dating back to the 1980s. Our counterfactual suggests that the pay caps more than account for this decline, with inequality increasing by 0.4pp under the counterfactual rather than falling by 1.9pp as observed.

In 2010, public sector workers earned a conditional wage premium of 13.7% over observationally equivalent private sector counterparts, but by 2019 this had fallen to 8.6%, significantly reducing the monetary returns to public sector employment. Given that women and workers outside the South of England are disproportionately represented in the public sector, we show that the pay caps had large and significant implications for gender and regional wage inequality, increasing the gender pay gap by 0.9pp and the regional gap by 0.4pp. Contemporaneous austerity policies, which reduced the public sector employment share from 30.8% to 23.8%, further exacerbated these disparities, increasing the gender and regional gaps to 1.9pp and 0.7pp. Relative to the observed baseline in 2019, this corresponds to a 17.5% and 9.9% increase, respectively.

Taken together, our results highlight the important role that the public sector as an institution plays in shaping overall inequality. The existing literature in this field has omitted the public sector in this way, instead focusing on other institutions such as the minimum wage or unionisation. Those that do study the public sector almost exclusively focus on the relative wage differential along the distribution. Our analysis also remains pertinent to the current policy debate: since 2021, the UK experienced its largest wave of public sector strikes since 1985, with 3 million working days lost—driven largely by disputes over pay. For example, the average doctor now earns 13% less in real terms than in 2010. Notably, more than half of OECD countries adopted public sector pay freezes after the financial crisis, suggesting that the dynamics we document extend well beyond the UK.

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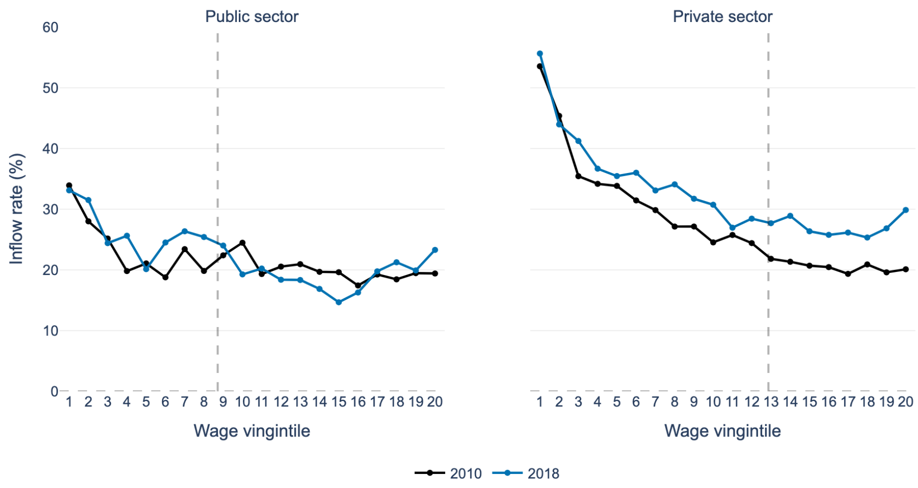
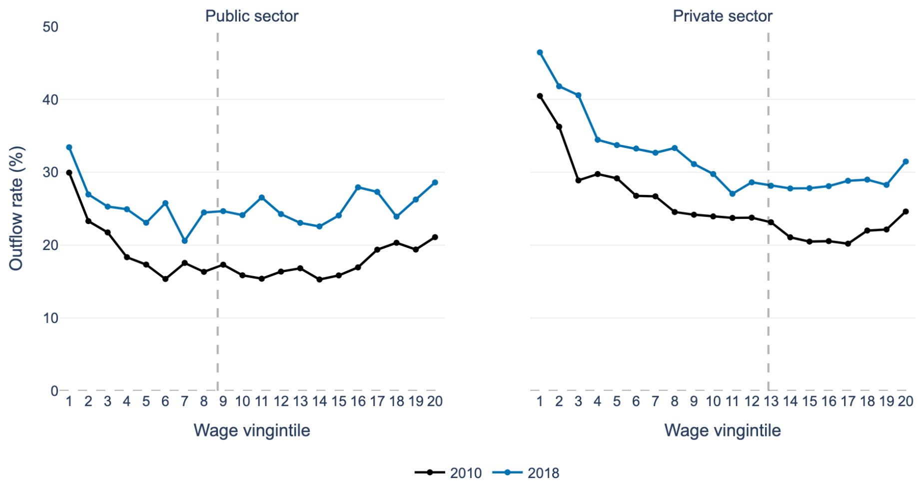
Appendix Figures & Tables

Figure A1: Smoothed Public Share of Employment Along the Wage Distribution

Note: The figure uses the unconditional wage distribution in 2010 and 2018 to compute the public sector share of employment across the wage distribution in 0.1 log-point intervals. Wages are measured hourly and the vertical line marks the pay cap threshold. *Source:* ASHE.



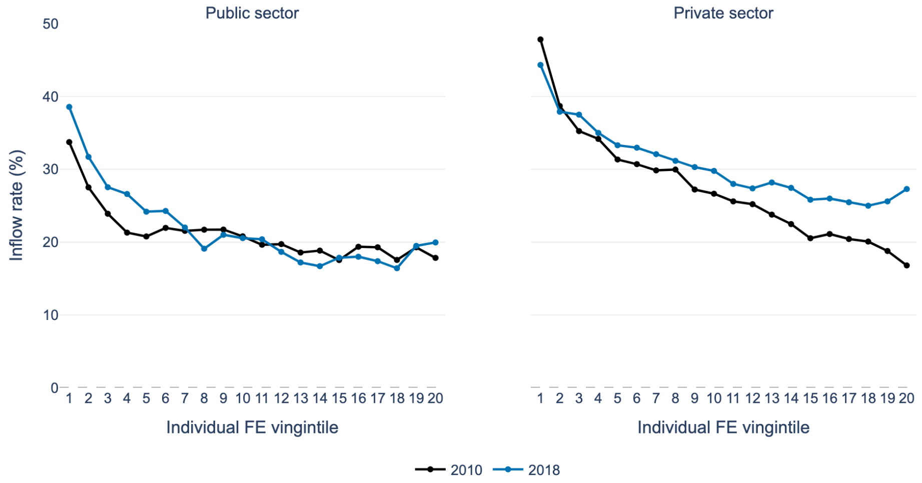
(a) Inflow



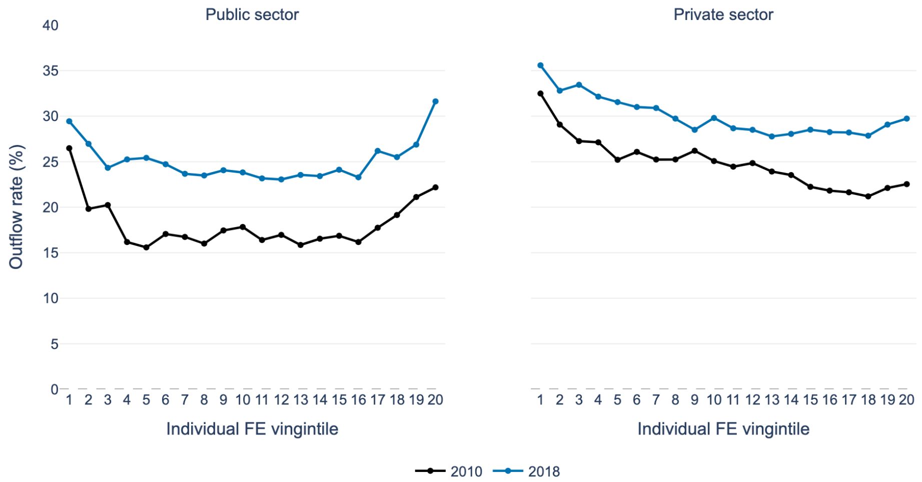
(b) Outflow

Figure A2: Inflow and Outflow Along the Wage Distribution

Note: Panel A shows inflow into each sector (from another sector) between 2009-2010 (black) and 2017-2018 (blue). Inflow is evaluated along 20 vigintiles of the wage distribution. Panel B shows outflow (either to another sector or to non-observation) along the same dimensions. *Source*: ASHE.



(a) Inflow



(b) Outflow

Figure A3: Inflow and Outflow Along the Individual Fixed Effect Distribution

Note: Panel A shows inflow into each sector (from another sector) between 2009-2010 (black) and 2017-2018 (blue). Inflow is evaluated along 20 vigintiles of the individual fixed effect distribution. Individual fixed effects are estimated using equation (1) without sector time interactions. Panel B shows outflow (either to another sector or to non-observation) along the same dimensions. *Source: ASHE.*

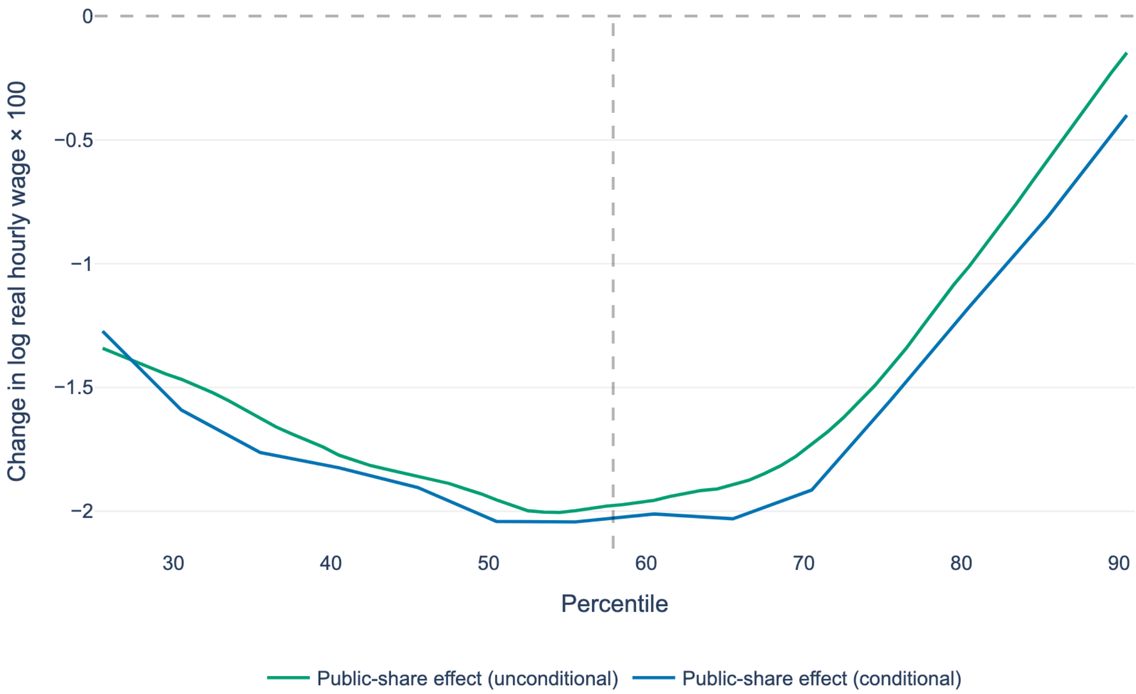


Figure A4: Conditional and Unconditional Public Share effects

The figure presents the public sector share effect, estimated using a DiNardo et al. (1996) decomposition. The conditional version (from the main text) controls for sex, age (squared), full-time status, region fixed effects, as well as a sectoral indicator. The unconditional version just includes a sectoral identifier. The figure also serves as a robustness check relating to Figure 7. *Sources:* ASHE.

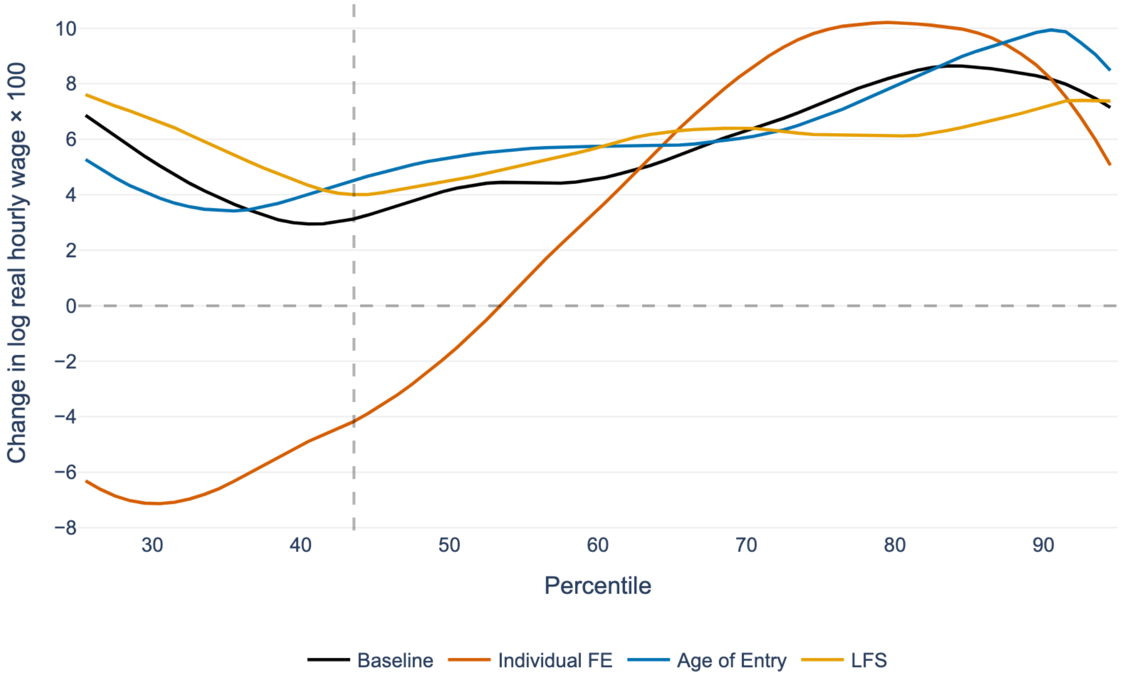
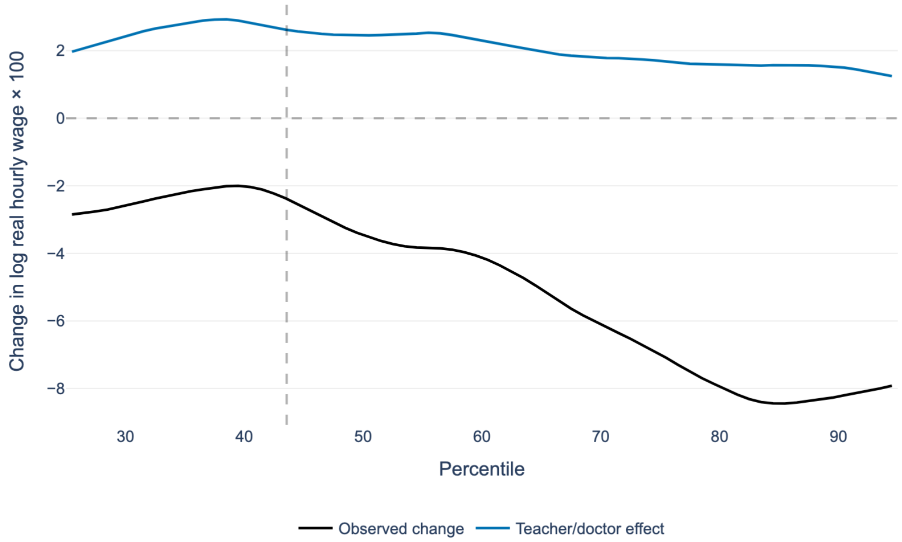
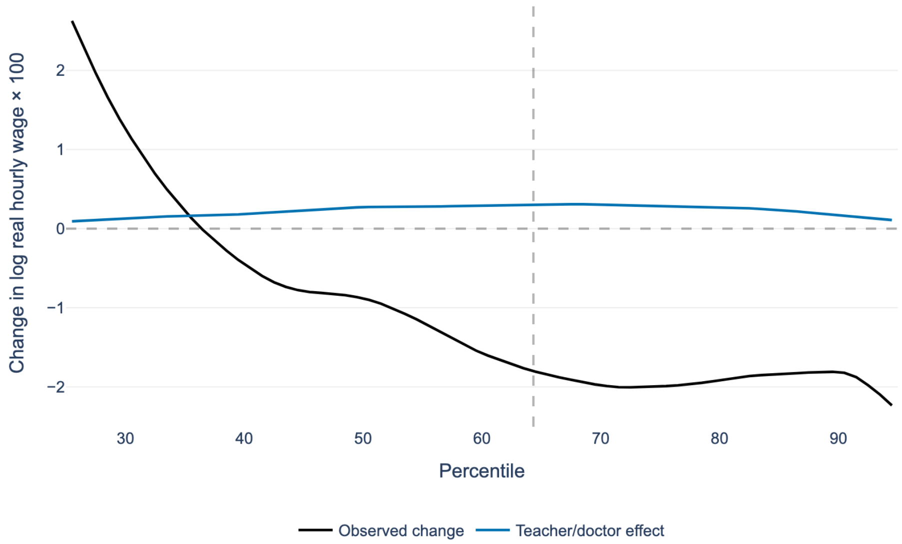


Figure A5: Public Sector Relative Price Effect Robustness Checks

The figure presents relative price effects in the public sector, estimated using a DiNardo et al. (1996) decomposition that controls for sex, age (squared), full-time status, and region fixed effects. The figure serves as a robustness check relating to Figure 5c. The first robustness check includes pre-period individual fixed effects in the decomposition, the second includes age of labour market entry cells and the third replicates the data using the LFS. *Sources:* ASHE and LFS.



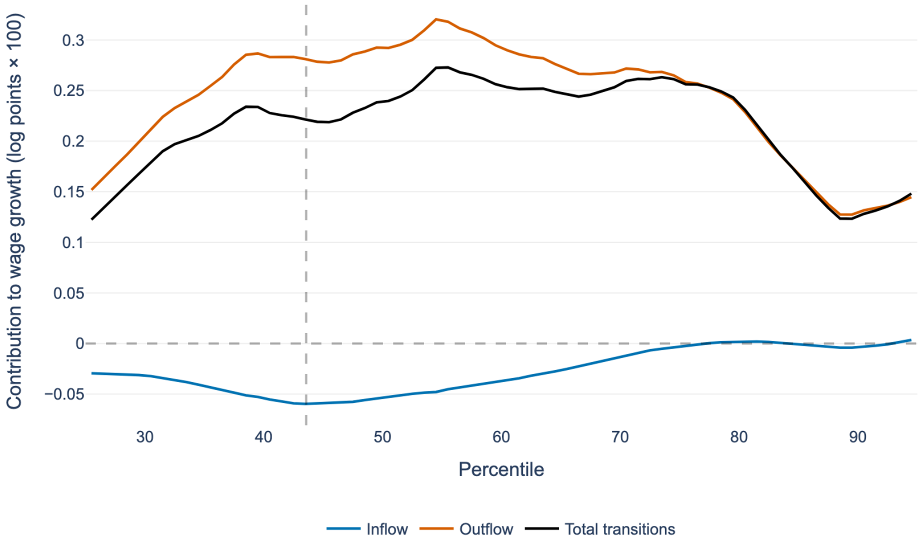
(a) Public



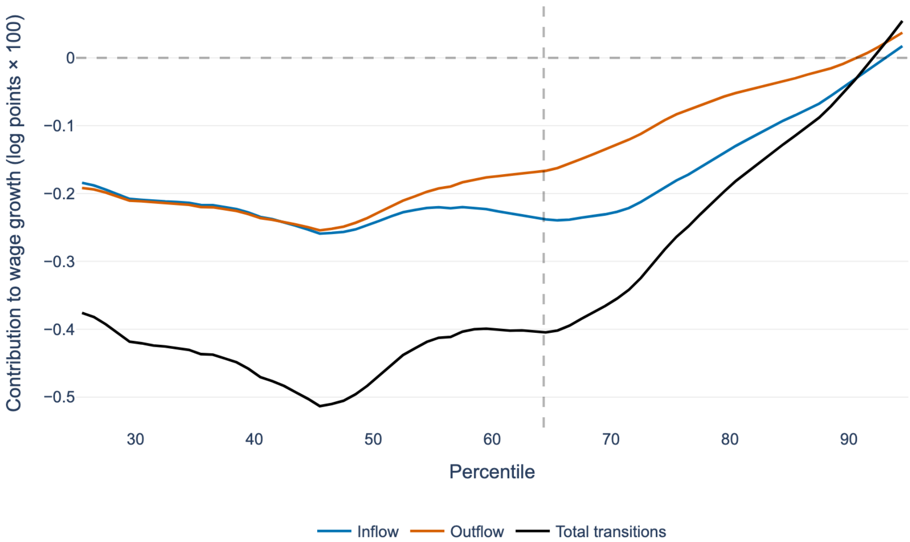
(b) Private

Figure A6: Teacher-Doctor Decompositions

Note: The figure presents characteristic effects in the public and private sectors, estimated using a DiNardo et al. (1996) decomposition that accounts for teachers and doctors. Each panel includes an overall estimate in black as well as the contribution of each characteristic. *Source:* ASHE.



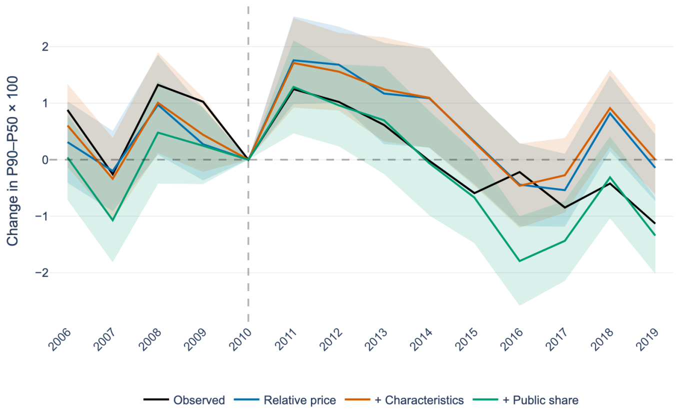
(a) Public



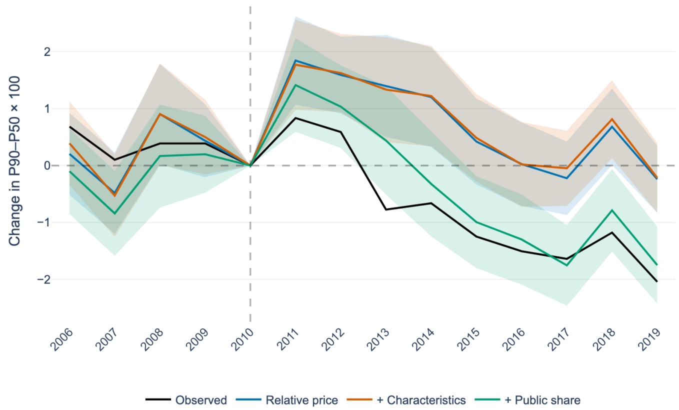
(b) Private

Figure A7: Transition Decompositions

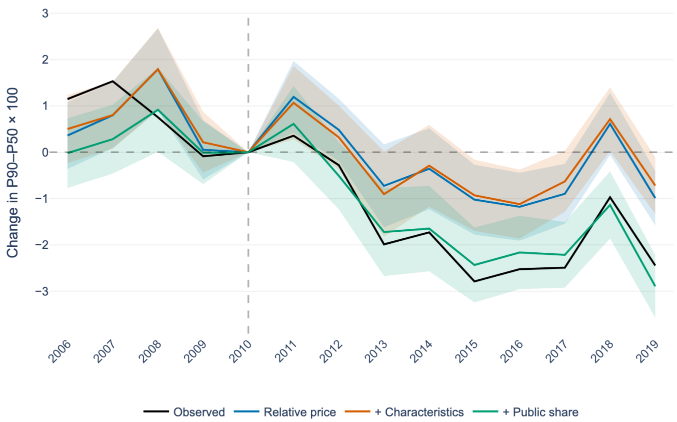
Note: The figure presents characteristic effects in the public and private sectors, estimated using a DiNardo et al. (1996) decomposition that accounts for inflow and outflow status. Each panel includes an overall estimate in black as well as the contribution of each characteristic. *Source:* ASHE.



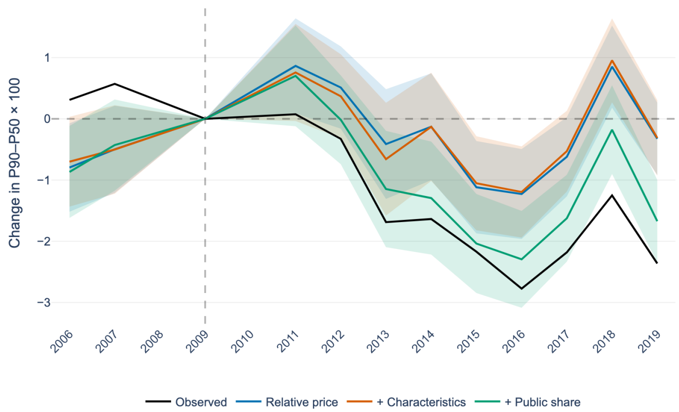
(a) Weekly



(b) Full time equivalent



(c) No overtime



(d) 2008-10 normalisation period

Figure A8: P90 P50 Robustness Checks.

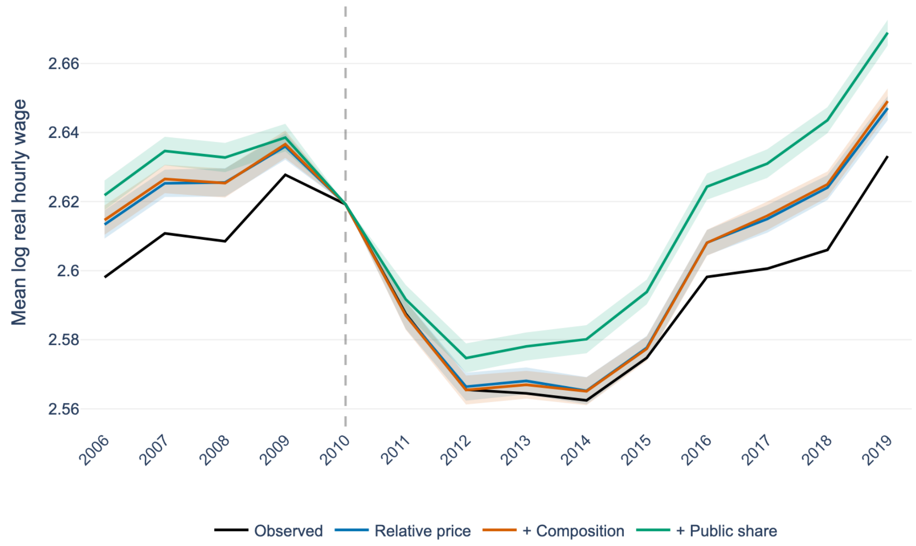
Note: Based on ASHE sample for public and private sector workers aged 16 and over. The figure displays observed and counterfactual inequality, measured by the difference between the 90th and 50th percentiles of the wage distribution. The robustness outcomes include weekly wages, full time equivalent wages, wages without overtime and hourly wages where 2008-10 is the base period. Counterfactuals are calculated using the decomposition method of DiNardo et al. (1996), which conditions on sex, age (squared), full time status, and region fixed effects. Standard errors are calculated following Chernozhukov et al. (2013).

(a) Female

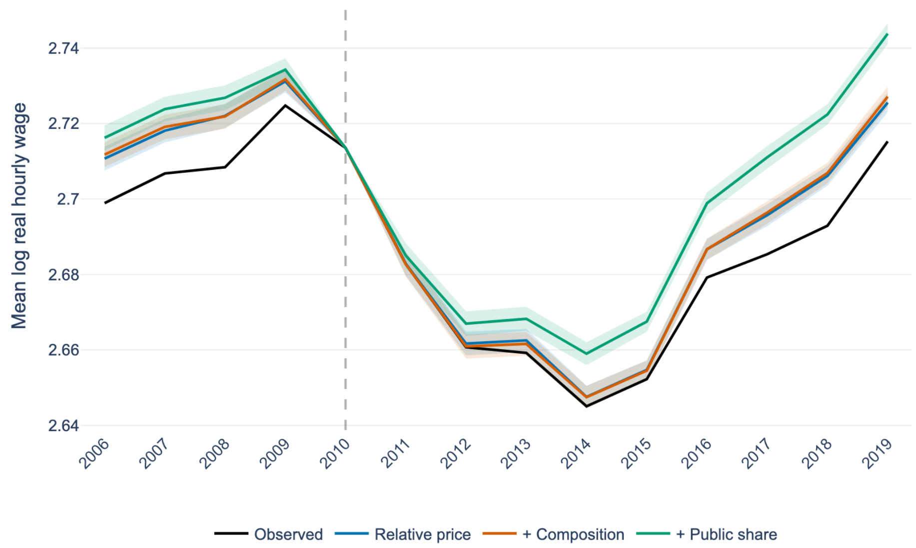
(b) Male

Figure A9: Counterfactual Gender Pay Gap

Note: The figure displays observed and counterfactual inequality, measured by the difference between the mean male and mean female wage. The difference between the black and blue lines represents the relative price effect, the wage distribution that would have prevailed had public sector returns to observable characteristics increased in line with those in the private sector. The difference between the blue and orange lines captures the effect of changes in public sector characteristics. The difference between the orange and green lines reflects the additional impact of holding the public sector employment share constant. Counterfactuals are calculated using the decomposition method of DiNardo et al. (1996), which conditions on sex, age (squared), full time status, and region fixed effects. Standard errors are calculated following Chernozhukov et al. (2013). *Source*: ASHE.



(a) North of England, Wales and Scotland



(b) South of England

Figure A10: Counterfactual Regional Wages

Note: The figure shows observed and counterfactual regional wages. The difference between the black and blue lines represents the relative price effect, the wage distribution that would have prevailed had public sector returns to observable characteristics increased in line with those in the private sector. The difference between the blue and orange lines captures the effect of changes in public sector characteristics. The difference between the orange and green lines reflects the additional impact of holding the public sector employment share constant. Counterfactuals are calculated using the decomposition method of DiNardo et al. (1996), which conditions on sex, age (squared), full time status, and region fixed effects. Standard errors are calculated following Chernozhukov et al. (2013). The South comprises of London, the South East, the South West, and the East of England; the rest of the UK covers all other government office regions.*Source*: ASHE.

Table A1: Public and Private Sector Employment by Two-Digit Occupation

Two-digit SOC 2010 occupation	Private workers	Public workers	Public sector share
11 Corporate managers and directors	202906	26289	11.1%
12 Other managers and proprietors	47136	9906	17.2%
21 Science, research, engineering and technology professionals	116919	20320	14.2%
22 Health professionals	24630	128346	83.3%
23 Teaching and educational professionals	11315	119305	90.7%
24 Business, media and public service professionals	47527	23585	31.8%
31 Science, engineering and technology associate professionals	58376	15986	21.0%
32 Health and social care associate professionals	8125	26984	76.8%
33 Protective service occupations	2489	49905	95.2%
34 Culture, media and sports occupations	19068	5013	19.7%
35 Business and public service associate professionals	192867	41603	16.9%
41 Administrative occupations	295505	168923	35.5%
42 Secretarial and related occupations	90407	20322	18.2%
51 Skilled agricultural and related trades	7040	3675	33.0%
52 Skilled metal, electrical and electronic trades	132688	5969	4.2%
53 Skilled construction and building trades	43624	5305	10.3%
54 Textiles, printing and other skilled trades	56755	5349	8.7%
61 Caring personal service occupations	90210	128695	58.0%
62 Leisure, travel and related personal service occupations	36562	14853	28.4%
71 Sales occupations	274960	3029	1.1%
72 Customer service occupations	70304	10309	12.6%
81 Process, plant and machine operatives	135617	4470	3.1%
82 Transport and mobile machine drivers and operatives	107754	7972	6.8%
91 Elementary trades and related occupations	68915	3104	4.2%
92 Elementary administration and service occupations	301868	97046	23.6%

Notes: The table reports the number of public- and private-sector worker observations in each two-digit Standard Occupational Classification (SOC) 2010 occupation. The public sector share is calculated as the number of public-sector worker observations divided by the total number of public- and private-sector worker observations within each occupation. The presence of observations from both sectors within each occupation demonstrates common support across the occupational categories used in the analysis. *Source*: ASHE.

Table A2: Change in conditional public-sector premium since 2010, by wage bin (log points × 100); standard errors in parentheses

Year	Bin 1	Bin 2	Bin 3	Bin 4	Bin 5	Bin 6	Bin 7	Bin 8	Bin 9	Bin 10
2006	0.18	-0.65	-0.39	-0.33	-0.20	-0.61	-0.90	-1.26	-0.88	-1.61
	(0.26)	(0.14)	(0.14)	(0.13)	(0.13)	(0.15)	(0.16)	(0.19)	(0.23)	(0.44)
2007	-0.49	-0.77	-0.28	-0.63	-0.62	-0.60	-0.97	-1.31	-1.22	-2.92
	(0.25)	(0.15)	(0.14)	(0.13)	(0.13)	(0.15)	(0.17)	(0.19)	(0.23)	(0.44)
2008	-1.26	-0.22	-0.04	-0.41	-0.50	-0.63	-0.97	-1.31	-1.37	-4.22
	(0.25)	(0.15)	(0.14)	(0.13)	(0.13)	(0.15)	(0.17)	(0.19)	(0.23)	(0.45)
2009	0.05	-0.24	-0.04	-0.32	-0.08	-0.34	-0.68	-0.41	-0.37	-1.59
	(0.25)	(0.14)	(0.14)	(0.12)	(0.13)	(0.14)	(0.16)	(0.19)	(0.22)	(0.43)
2010	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
2011	-0.48	-0.09	0.02	0.31	-0.21	-0.21	-0.43	-0.14	-0.10	-0.69
	(0.27)	(0.15)	(0.13)	(0.12)	(0.12)	(0.14)	(0.16)	(0.18)	(0.22)	(0.43)
2012	-1.07	-0.22	-0.21	0.19	-0.45	-0.39	-0.68	-0.54	-0.61	-1.37
	(0.28)	(0.15)	(0.13)	(0.12)	(0.12)	(0.14)	(0.16)	(0.18)	(0.22)	(0.43)
2013	-1.07	-0.38	-0.36	0.24	-0.10	-0.39	-0.99	-0.83	-1.06	-2.33
	(0.28)	(0.15)	(0.13)	(0.12)	(0.12)	(0.14)	(0.16)	(0.18)	(0.22)	(0.43)
2014	-0.98	-0.37	-0.39	0.18	-0.02	-0.49	-1.13	-1.05	-1.36	-2.68
	(0.29)	(0.15)	(0.13)	(0.12)	(0.13)	(0.14)	(0.16)	(0.18)	(0.22)	(0.43)
2015	-0.26	-0.22	-0.26	-0.00	-0.31	-0.79	-1.23	-1.01	-1.81	-3.60
	(0.43)	(0.15)	(0.13)	(0.12)	(0.13)	(0.15)	(0.16)	(0.19)	(0.22)	(0.43)
2016	-1.74	-0.42	-0.46	-0.48	-0.24	-0.85	-1.42	-1.17	-2.23	-4.59
	(0.39)	(0.14)	(0.13)	(0.12)	(0.13)	(0.15)	(0.16)	(0.19)	(0.23)	(0.44)
2017	-2.17	-0.81	-0.11	-0.27	-0.44	-0.92	-1.57	-1.36	-2.66	-5.51
	(0.43)	(0.14)	(0.13)	(0.12)	(0.13)	(0.15)	(0.16)	(0.19)	(0.23)	(0.44)
2018	-2.22	-0.81	-0.69	-0.81	-1.25	-1.11	-2.06	-1.70	-3.32	-6.89
	(0.31)	(0.14)	(0.13)	(0.12)	(0.13)	(0.15)	(0.17)	(0.19)	(0.23)	(0.45)
2019	-2.33	-0.04	0.07	-0.01	-0.45	-0.72	-1.03	-0.93	-2.37	-6.49
	(0.40)	(0.14)	(0.13)	(0.13)	(0.13)	(0.15)	(0.17)	(0.19)	(0.23)	(0.45)

Notes: The table reports the conditional wage premium earned by public-sector workers relative to private-sector workers, relative to 2010, across 10 deciles of the residual wage distribution. Wages are residualised with respect to age, age squared, union status, and full-time status, as well as individual and time fixed effects. *Source*: ASHE.

Table A3: Summary of Decomposition Effects

(1) Effect	(2) Sub-Effect	(3) Definition	(4) Factor Held Constant	(5) Sample
Characteristic	-	Change in wage distribution due to changes in the distribution of observable worker characteristics	Returns (β)	Public only
Wage Structure	-	Change in wage distribution due to evolving returns to observable characteristics between 2010 and τ	Composition (x)	Public only
	Relative Price	Change in the wage distribution due to changing returns to observable characteristics in the public sector relative to the private sector	Composition (x)	Public only
	Remainder	Change in the wage distribution due to public sector wage structure changes not explained by the private sector counterfactual	Composition (x)	Public only
Share	-	Change in overall wage distribution due to falling public sector share of employment	Returns (β), Composition (x)	Public and private sector

Note: Table of counterfactual effects identified throughout Section 4. Column (1) defines the effect, and column (2) identifies any sub-effects within this group. Column (3) explains the effect; column (4) outlines the component of the distribution held fixed. Column (5) states the data used to identify the effect.

A Causal Effects of the Pay Caps on Earnings

To assess whether the pay caps caused the decline in public sector wages documented in the main text, and whether private sector wage setting responded to the policy, we estimate a triple-difference event-study specification. The first difference compares workers employed in the public and private sectors in 2010, treating the private sector as the control group, as in the counterfactual analysis. The second compares workers above and below the £21,000 pay-cap threshold based on their 2010 earnings. The third compares the evolution of this capped–uncapped wage differential across sectors over time.

Our preferred specification restricts the sample to workers earning between £15,000 and £35,000 in 2010. This provides a relatively large treatment and control sample while keeping workers reasonably close to the policy threshold. We avoid extending the lower bound further because of the minimum wage, which corresponded to annual earnings of approximately £9,997 for an individual working 33 hours per week in 2010. We include individual and year fixed effects and cluster standard errors at the individual level. Our specification is:

$$
\begin{aligned}
y_{it} = \alpha_i + \gamma_t + &\sum_{\tau \neq 2010} \delta_\tau \left(\text{Public}_{i,2010} \times \mathbb{1}\{t = \tau\}\right) + \sum_{\tau \neq 2010} \lambda_\tau \left(\text{Capped}_{i,2010} \times \mathbb{1}\{t = \tau\}\right) + \\
&\sum_{\tau \neq 2010} \beta_\tau \left(\text{Public}_{i,2010} \times \text{Capped}_{i,2010} \times \mathbb{1}\{t = \tau\}\right) + \varepsilon_{it}.
\end{aligned}
\tag{12}
$$

where y_{it} is the outcome for individual i in year t, α_i and γ_t denote individual and year fixed effects, respectively, $\text{Public}_{i,2010}$ identifies workers employed in the public sector in 2010, and $\text{Capped}_{i,2010}$ identifies workers whose 2010 earnings exceeded the pay-cap threshold. We define an individual as capped if their full time equivalised income (based on a 33 hour per week threshold) exceeds the cap value. The omitted year is 2010. The δ_τ coefficients allow public and private sector outcomes to evolve differently over time irrespective of cap status, while the λ_τ coefficients allow workers above and below the threshold to follow different time paths irrespective of sector. The coefficients of interest, β_τ, therefore measure the change in the capped–uncapped outcome gap in the public sector relative to the corresponding change in the private sector.

Identification requires that, absent the pay caps, the capped–uncapped outcome differential would have evolved similarly across the public and private sectors. We assess the plausibility of this assumption using the pre-policy coefficients: differential pre-trends would appear as non-zero estimates of β_τ before 2010.

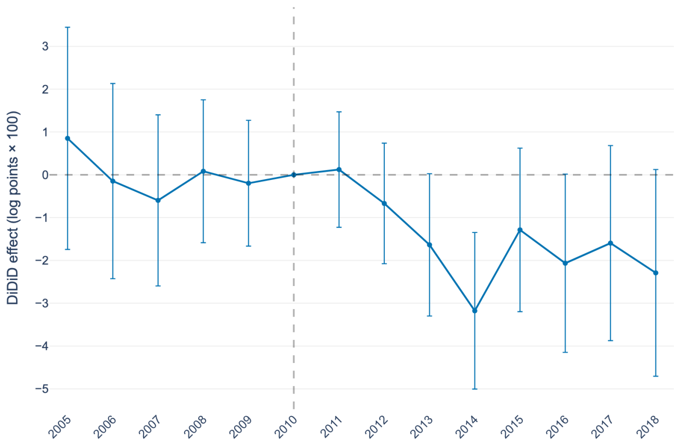
Because the triple difference identifies only the relative divergence between sectors, it does not reveal whether that divergence originates in the public sector, the private sector, or both. We therefore also estimate separate event-study difference-in-differences specifications within each sector. These compare the evolution of outcomes for workers above and below the 2010 threshold:

$$y_{it} = \alpha_i + \gamma_t + \sum_{\tau \neq 2010} \beta_\tau^s \left(\text{Capped}_{i,2010} \times \mathbb{1}\{t = \tau\}\right) + \varepsilon_{it}, \qquad s \in \text{Public}, \text{Private}. \tag{13}$$

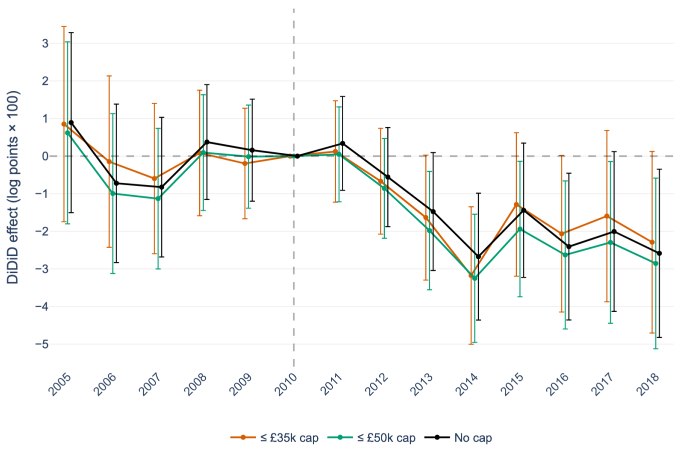
We use log real hourly wages as the outcome and restrict the analysis to workers who remain in their 2010 sector throughout the sample. This allows us to focus on wage adjustment among continuously exposed workers rather than allowing sectoral mobility to attenuate the estimated effect of the caps. Figure A11a shows that, between 2010 and 2015, the wage gap between capped and uncapped workers declined by approximately 3% more in the public sector than in the private sector, with the effect persisting through the end of the sample.

Figure A11b relaxes the upper earnings restriction to examine whether the result extends further up the wage distribution. We consider samples capped at £50,000 and with no upper earnings restriction. The estimates are very similar across specifications, with relative public sector wages falling by approximately 3% by 2014. The coefficients also become modestly more negative as the upper earnings restriction is relaxed, consistent with the regressive pattern documented in the distributional analysis. Because these broader samples still include the large mass of workers earning just above the threshold, the point estimates represent a pooled average heavily weighted toward those with lower counterfactual earnings. This explains why the average event-study impact of approximately 3% is smaller than the 7.6 percentage-point effect isolated at the public-sector 90th percentile, where the flat nominal cap binds much more tightly on counterfactual wage growth. We stop short of attributing causality to the additional specifications in this second panel as the composition of control and treatment groups begin to diverge when the distance from the cutoff increases. However, the lack of pre-trend in any specification provides very strong suggestive evidence that we are indeed capturing the caps.

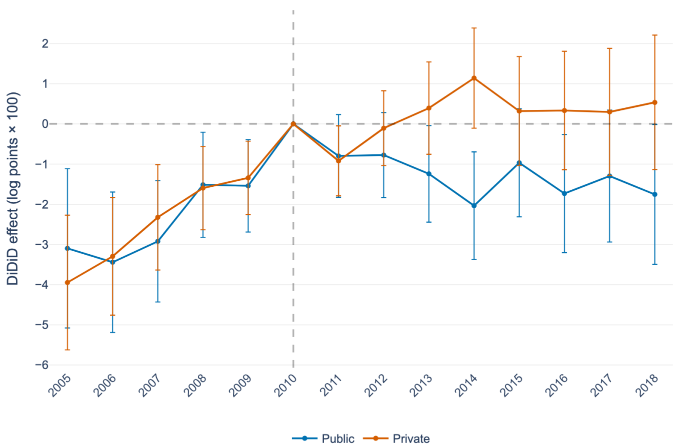
Figure A11c plots the sector-specific difference-in-differences estimates. Although the capped–uncapped wage gap widened within both sectors before 2010, the two series evolved similarly, producing little evidence of a differential pre-trend in the triple difference. After 2010, the capped–uncapped wage gap declines in the public sector relative to the private sector. In the private sector, by contrast, the gap continues to widen at approximately its pre-policy rate until 2015 before flattening somewhat. The evidence therefore indicates that the triple-difference effect is driven primarily by a deterioration in the relative wages of capped public sector workers rather than by a contemporaneous change in the relative wage trajectory of comparable private sector workers. We consequently find little evidence that private sector wage setting responded materially to the caps, supporting its use as the counterfactual wage structure in the distributional analysis.



(a) Triple Difference



(b) Triple Difference with Various Treated Groups



(c) Difference in Difference

Figure A11: Wage Event Studies

Note: The figure shows triple difference and difference in difference estimates from equations (12) and (13) where the outcome is log real hourly wage. *Source*: ASHE.