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Received: 9 January 2023 | Accepted: 24 October 2023

DOI: 10.1111/bjir.12778

BJIR An International Journal of Employment Relations LSE

ORIGINAL ARTICLE

Pay transparency intervention and the gender pay gap: Evidence from research-intensive universities in the UK

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All statistics in this paper follow a level of aggregation to maintain the anonymity of individuals and ensures no personal data or personally sensitive data are identifiable. We follow the Higher Education Statistic Agency (HESA) standard rounding methodology to comply with HESA agreement. Rounding method implies that (i) counts of individuals are rounded to the nearest multiple of 5; (ii) percentages based on fewer than 22.5 individuals are suppressed; and (iii) averages based on seven or fewer individuals are suppressed.

Abstract

This study investigates the impact of a pay transparency intervention in reducing the gender pay gap in the UK university sector. Introduced in 2007, the initiative enabled public access to average annual earnings disaggregated by gender in UK universities. We use a detailed matched employee-employer administrative dataset that follows individuals over time, allowing us to adopt a quasi-experimental approach based on event studies around the intervention. We find that the earnings of female academics increased by around 0.62 percentage points compared to their male counterparts as the control group, whose earnings remained constant after the pay transparency intervention, reducing the gender pay gap by 4.37 per cent. Further evidence suggests that the main mechanism for the fall in the pay gap is driven by female employees negotiating higher wages, particularly among senior female academics.

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1 INTRODUCTION

The gender pay gap, although narrowing over time, remains a persistent problem in labour markets globally, with women earning less than men on average, as discussed in recent overviews by Goldin (2014), Blau and Kahn (2017) and Cozzi et al. (2018). In fact, the pay gap in the higher education sector (10.8 per cent) is higher than the national average for full-time employees (7.9 per cent) (White, 2021). Findings from representative surveys and administrative records indicate that the gender pay gap in universities is large (Gamage et al., 2020; Mumford & Sechel, 2020) and that progress in addressing the issue is slow. The quest for better and more efficient policies to combat the problem is ongoing. This article exploits an information intervention resulting from a public data release on academic earnings to provide evidence on the causal impact of pay transparency regulation on the gender pay gap in the UK higher education sector.

Pay transparency interventions are becoming increasingly prominent in efforts to tackle the gender pay gap. This is based on the assumption that the gap persists partly because it is 'hidden', which has led organizations, such as the Fawcett Society (2018), to believe that the culture of pay secrecy facilitates pay discrimination to thrive.[1] At the policy level, in 2006, the Danish government mandated firms with more than 35 employees to report wages by gender (Bennedsen et al., 2022); Australia, Canada, France and Germany have introduced some elements of pay reporting/transparency to their equal pay legislation, such as publishing average male and female pay or scores of gender inequality indicators.[2] In the UK, the 2010 Equality Act (Gender Pay Gap Information) mandated employers with over 250 employees to publish gender pay gap data in 2017 (Abudy et al., 2023; Duchini et al., 2022a; Jones & Kaya, 2023).[3]

There are several underlying economic theories that can explain how pay transparency may decrease the gender pay gap. Under a simple Beckerian or statistical model of discrimination, incomplete information – that is not knowing what others earn — may result in discriminatory practices going undetected in the long run (Kerwin & Guryan, 2007). For example, any lack of information about peers' wages can explain some of the evidence showing how women are less likely to ask for a pay rise or to negotiate their salary when offered a job (Babcock et al., 2003; Bowles et al., 2007; Cullen & Pakzad-Hurson, 2023; Leibbrandt & List, 2015).[4] There is indeed causal evidence that wage transparency in the presence of gender gaps leads women to ask for higher pay, eliminating any discrepancy (Roussille, 2022). Another theory builds on the reputational effects that firms may incur from paying a lower wage to women once information is released. Duchini et al. (2022a) show that there is a significant correlation between the publication of gender pay equality indicators and the reputation that companies hold among consumers and investors.

In this study, we adopt a difference-in-differences like approach based on sharp changes around a gender pay publication by the Times Higher Education guide ('*THE*' henceforth).[5] In 2007, *THE* started publishing average annual earnings data by gender and academic rank, covering about 99 per cent of universities. There was no formal announcement prior to publication, so universities had no room to reduce the gender pay gap in advance of the publication.[6] We use administrative data from the Higher Education Statistical Agency (HESA) containing earnings of academics employed in research-intensive institutions between 2004 and 2016. The panel nature of the data allows us to compare within-individual changes, before and after the pay transparency initiative, for women and men, after controlling for individual-level time variant and invariant characteristics, time fixed effects, university-specific fixed effects and university-specific time fixed effects. The smoothness of male outcomes around the policy intervention supports the use of men as a 'control group'.

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We find a 4.37 per cent decrease in the gender pay gap following the pay transparency intervention. This reduction is driven by a 0.62 percentage point increase in female earnings, while male earnings remain constant. Our results are robust to considering a balanced sample of individuals who are observed over the entire period: female earnings increase by 1.27 percentage points, reducing the pay gap by 11.59 per cent. We exploit the big sample sizes in the administrative data to run heterogeneity analysis alongside several dimensions to shed light onto the mechanisms at play. We find that our results are driven by women at the top of the earnings distribution who experience a reduction in the gender pay gap three times greater than women at the bottom of the earnings distribution following the intervention. Exploiting the longitudinal nature of our data allows us to uncover that the impact of pay reporting likely comes from women progressing within their institution, rather than from women moving between institutions.

A key challenge for our identification is whether the pay transparency intervention may have been correlated with pre-existing pay gap trends. We address this issue in several ways. First, we use a difference-in-differences like event study methodology that exploits the fact that the pay transparency shock generated sharp changes in the gender pay gap which are arguably orthogonal to any unobserved determinants of the wage gap that evolve more smoothly. Under this assumption, our approach can successfully identify the causal effect of the 2007 information intervention for female workers relative to male workers. We find that while men's and women's earnings evolve in a parallel fashion before the intervention kicks in, they start to converge immediately afterwards. Second, we show that, after controlling for the same characteristics we use in our benchmark regression, the year of the *THE* publication is not correlated with the initial gender pay gap levels, confirming that the information intervention induced by the publication is orthogonal to the prevailing pay gap. We also rule out the possibility that our findings may be driven by other policies introduced around the study period, such as the Athena SWAN intervention, aimed to increase female representation and promotion of women to higher academic ranks in the sciences discipline.

Our article contributes first and foremost to the literature on gender inequality in the labour market: for example, Altonji and Blank (1999), Bertrand (2010), Blau and Kahn (2017) and Olivetti and Petrongolo (2016).[7] Much of this literature has focused on the role of human capital, occupation and discrimination in explaining gender gaps and the role of parenthood (Kleven et al., 2019). A limited number of recent papers try to assess the causal role of information asymmetries on gender pay gaps using quasi-experimental approaches similar to ours (see Table 1).[8] The majority of studies find a reduction in the gender pay gap after introducing pay transparency measures. Our results are in line with these findings. In particular, UK-based studies find that gender pay gap falls by around 2.3–2.8 percentage points; three times the magnitude of our findings. Arguably, the highly regulated pay structure and the relatively low level of awareness of the policy among academics are plausible explanations of the low impact of this intervention in the academic sector (Pfefer, 2020).

The large sample sizes of the HESA admin data, its longitudinal nature and the idiosyncrasy of the UK Higher Education Sector allow us to uncover previously unexplored theoretical and empirical mechanisms behind the reduction in the pay gap. The literature tends to document a progressive reduction in the gender pay gap after the introduction of a pay transparency intervention. Bennedsen et al. (2022) analyse the Denmark framework for reporting gender-based disaggregated statistics and find no effect on female wages but a 3.8 per cent drop in male salaries. Similarly, Baker et al. (2023) look at the effects of the 1996 Disclosure Act in Canada and find a drop in male salaries relative to the control group post-reform. Lastly, Duchini et al. (2022a) document

TABLE 1 Previous studies on the effect of pay information shock and the gender pay gap.

Policy	Country	Authors and year	Publication	Data source	Data years	Method	Overall effect on pay gap	Effect on women's wages	Effect on men's wages
2004 ServiCo measures to improve pay transparency	US	Castilla (2015)	Organization Science	ServiCo employees	1996–2003 2005–2009	Cross-sectional time-series linear models using the method of generalized estimating equations (GEE) with individual pay as outcome variable	Pay transparency reduces gender pay gap	N/A	N/A
State bans on pay secrecy	US	Kim (2015)	Industrial Relations	Current Population Survey	1977–2012	Difference in differences comparing wages of men and women in states that outlawed pay secrecy	4.8–4 pp fall (Causal)	Increased 4–5%	N/A
Local Government Compensation Reporting Program, 2010	US	Mas (2017)	Journal of Political Economy	CalPERS data multiple records	2000–2012	Difference in differences comparing wages of men and women in new disclosure cities following the intervention.	N/A [(b)]	N/A	N/A
State bans on pay secrecy	US	Burn and Kettler (2019)	Labour Economics	Current Population Survey	1977–2016	Difference in differences comparing wages of male and female managers	No effect	<median: increase of wages of 2.9%	<median: small, but insignificant, decline in wages & >median: increase of 2.7%
National Action Plan for gender equality in the labour market, 2010	Austria	Böheim and Gust (2021)	IZA WP	Austrian SocialSecurity Database(ASSD)	2009–2017	Regression Discontinuity Design exploiting the cutoff of firm size at 1000 employees above which the transparency reform applies	Wage transparency law did not reduce the gender wage gap	For newly hired women in larger firms, the law resulted in a wage increase	N/A
Pay Transparency Act, 2017	Germany	Ahrens and Scheele (2022)	German Politics	N/A	N/A	Qualitative assessment	N/A	N/A	N/A

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TABLE 1 (Continued)

Policy	Country	Authors and year	Publication	Data source	Data years	Method	Overall effect on pay gap	Effect on women's wages	Effect on men's wages
Act No. 562: requires firms to report gender-based disaggregated stats	Denmark	Bennedsen et al. (2022)	Journal of Finance	Integrated Database for Labor Market Research	2003–2008	Difference in differences comparing wages of men and women in treated firms following the intervention	2 pp fall (Causal)	No significant effect of regulation	3.8-percentage-point-larger drop in the wage growth
Equality Act 2010 (Gender Pay Gap Information) Regulations 2017	UK	Duchini et al. (2022a)	SSRN	Annual Survey of Hours & Earnings (ASHE)	2013–2019	Difference in differences comparing wages of the treated and controlled firms after the intervention separated by gender	18% reduction in gender pay gap (Causal)	No significant effect	2.6% decrease in men's real hourly pay
Gender pay transparency policies on labour market	Denmark, Austria, UK and Canada	Duchini et al. (2022b)	WP	N/A	N/A	Literature review	Transparency policies can reduce the gender pay gap with limited costs for firms, but may not be suited to achieve the objective of improving outcomes for lower-paid employees	N/A	N/A
Times Higher Education Pay reporting	UK	Gamage et al. (2022)	SSRN	Higher Education Statistical Agency administrative data	2004–2016	Difference in differences comparing wages of males and females before and after the intervention	4.37% decrease in gender pay gap following pay transparency intervention	0.62 pp increase	Remain constant

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TABLE 1 (Continued)

Policy	Country	Authors and year	Publication	Data source	Data years	Method	Overall effect on pay gap	Effect on women's wages	Effect on men's wages
Staggered shocks to the accessibility of information on wages in public university systems	US	Obloj and Zenger (2022)	Nature Human Behaviour	About 100,000 US academics	1997–2017	Static difference in differences explaining changes to institution-department wage variance, where the treatment is the implementation of pay transparency	Post-intervention, average conditional gender pay gap narrowed, but remained substantial at 2.6%	N/A	N/A
Publishing mean salary on a job portal	US	Roussille (2022)	WP	Hired.com data	N/A [(c)]	Difference in differences comparing male and female ask wage before and after the intervention	3.9-3.4 fall (gender ask gap)	Ask for 3.2% more	Continue asking for roughly the same as they would have otherwise
Public Sector Salary Disclosure Act, 1996	Canada	Baker et al. (2023)	AEJ: AE	Statistics Canada's University and College Academic Staff System (UCASS)	1970–2017	Difference in differences comparing wages of men and women following the intervention and an event study to determine short-run and long-run impact	2.2–2.4 pp (Causal)	Institution and department fixed effects: Womens' salaries rise relative to the control groups post-reform. Individual fixed effects: Womens' salaries either do not change or fall	Institution and department fixed effects: Mens' salaries rise relative to the control groups post-reform. Individual fixed effects: Mens' salaries fall relative to control groups post-reform
Gender Pay Transparency policies within firms	Various	Bennedsen et al. (2023)	Journal of Economic Surveys	N/A	N/A	Literature review	Several recent studies show that increased transparency causally reduces the gender wage gap	N/A	N/A

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TABLE 1 (Continued)

Policy	Country	Authors and year	Publication	Data source	Data years	Method	Overall effect on pay gap	Effect on women's wages	Effect on men's wages
"Right of Workers to Talk" (ROWTT) laws	US	Cullen and Pakzad-Hurson (2023)	Econometrica	American Community Survey (ACS)	2000–2016	Theoretical bargaining model and a meta-analysis of the literature	N/A	N/A	Men's wages decline more than women's wages when evaluating the effect of pay transparency on within-firm wages [(a)]
Pay Transparency Law	Austria	Gulyas et al. (2023)	AEJ: EP	Austrian social security administration	1997–2018	Event study comparing gender wage gap in treated and controlled firms following the intervention	No effect	N/A	N/A

Note: Studies listed in chronological order of publication.

[(a)] Cullen and Pakzad-Hurson (2023) test their theoretical model with an event-study analysis of US state-level laws protecting the right of private sector workers to communicate salary information with their coworkers. Consistent with theoretical predictions, transparency laws decrease wages by approximately 2%; which are the smaller in magnitude when workers have low individual bargaining power. The conclusion from their meta-analysis is that: 'six of nine studies find that wages for men decline more than wages for women when evaluating the effect of pay transparency on within-firm wages, consistent with our model's predictions of wage compression when men's outside options are higher than women's' (p.768), which we include here on the effect for men.

[(b)] In Mas (2017), the main finding suggests that among top managers, disclosure led to an average decline in compensation of 7% and a 75% increase in their quit rate, relative to managers in cities where salaries are already disclosed. This finding is not included in the table (i.e. N/A) as the focus of the article is on how pay disclosure changes wage setting at the top of the distribution and not on how the gender pay gap varies as a result of transparency, which is the objective of the review provided in this table.

[(c)] As stated by Roussille (2022), 'the exact start and end dates of the period for which the data was made available cannot be disclosed, to preclude inference about market shares' (p.10).

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that pay disclosure policies do not affect female hourly salaries but lead to a drop in male wages looking at data from the Annual Survey of Hours and Earnings in the UK. Our results are closer to those of Burn and Kettler (2019) who, similar to our findings, report women salaries increasing in some part of the wage distribution. We leverage on the longitudinal nature, which allows us to track individuals over time, to document that the increases in female wages come about as a result of women moving up within their institution, particularly for female full professors. Overall, our results are consistent with a bargaining mechanism as found in the literature (Cullen & Pakzad-Hurson, 2023): whereas, pay transparency limits the bargaining power of employees under collective bargaining (i.e. as in the case of non-full Professors in the UK), it may not necessarily harm — and, in fact, might benefit — those who are more likely to be subject to individual bargaining (i.e. the case of full Professors in the UK).

The article is organized as follows. Section 2 outlines the institutional background; Section 3 describes the data and empirical approach; Section 4 presents the results; Section 5 explores possible mechanisms behind the pay transparency; and Section 6 concludes.

2 | INSTITUTIONAL AND INTERVENTION SETTING

2.1 | Pay structure and bargaining strategy

In 2004, the UK moved away from a nationally determined pay scale to a more flexible pay structure implemented by the UCU Framework Agreement.[9] The new structure was designed to promote pay and career progression in order to attract and retain academic staff. The convention introduced two distinct divisions in the pay structure of academics according to which below-full-professorial pay (henceforth referred to as 'Non-Professorial', including, for example, Lecturers, Senior Lecturers and Readers; the equivalent of Assistant Professors and Associate Professors in the United States) is determined by sector-wide collective bargaining between the university and UCU, and full-professorial (henceforth simply referred to as 'Professorial') pay is determined between the academic and the university via individual bargaining.

Non-professorial pay is governed by a negotiated fixed pay structure, known as the pay spine system, which consists of 51 spinal points that correspond to a particular wage. Multiple pay spines form the pay scale for a specific academic grade; for example, 'Grade 6' relates to a position encompassing assisting teaching or research activity (corresponding to the pay spine scale between 22 and 32, which translates into a wage between £19,068 and £25,626 based on the 2003/2004 pay spine point system). Depending on qualifications and years of experience, academics are placed onto a particular spinal point. Each year, the post holder moves up by one spinal point until they reach the maximum automatic increment point. The individual can apply for promotion from one grade to the next either on reaching the top of a pay scale or before. The promotion decision is made at the discretion of the promotion panel. Although the union recommends the structure, the pay structures for job ranks across universities can vary.

Professorial pay, on the other hand, is not determined by the recommended pay structure.[10] Generally, professorial pay is above the highest spine point and is individually negotiated between the university and the staff member. Higher education providers set their pay at the beginning of the academic year (usually in August or September), making it difficult for wages to adjust immediately. Hence, any wage increment including the automatic increase and 'contribution' pay will come to effect at the beginning of the next period (i.e. the following August/September).

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Increments on professorial pay are linked to a performance-based review, allowing professors to renegotiate pay for the next academic year.

Given the above distinct pay structures between professors and non-professors, the flexibility and associated strategy for bargaining for a higher salary can differ significantly. The tightly regulated pay structure leaves little room for non-professors to bargain. The set formula assigns the individual to a particular pay spinal point. However, a non-professor in some circumstance can negotiate to be on a higher pay spine point but this would be within the grade pay scale. On the contrary, professors are not bounded by a wage ceiling and, therefore, have some flexibility to negotiate higher wages. In this article, we argue that professors are at a better position to negotiate relative to non-professors in light of the pay transparency initiative and we explore how pay transparency affects the salaries of female professors and non-professors.

2.2 | Times higher education (*THE*) pay initiative

The *Times Higher Education* (*THE*) was established in 1971 as part of *The Times* newspaper in England. It was later re-launched as an independent publication in 2008. *THE* provides information about the global higher education sector and, as part of its activities, it reports news and issues related to the higher education sector in the UK. In 2007, *THE* began publishing university-level pay data to inform the public on gender wage disparities.

The first *THE* pay publication reported the average nominal pay of male and female full-time academics for the 2005/2006 academic year at the university level.[11] By default, the *THE* report publishes all university pay data; a university can, in principle, request that the *THE* refrains from publishing their data. The pay information for reporting is collected from the Higher Education Statistical Agency (HESA).[12]

The structure of the pay report has been amended several times. Whereas the first year of pay publication is aggregated at all academic ranks, in the following years (2008 and 2009), the pay data is disaggregated by broad academic ranks, namely Professor, Senior Lecturer (Associate Professor)/Researchers, Lecturer (Assistant Professor), researcher, other grades and all grades. From 2009 onwards, the pay reporting is only disaggregated by Professors and non-Professors and is available for full-time academic staff only. Although not explicitly stated in the reports, disaggregation of pay data by Professor and non-Professors may be influenced by the differences in wage determination, which we discuss in our analysis.

3 | DATA AND EMPIRICAL STRATEGY

The administrative data for this study come from HESA. HESA acts as the intermediary between the higher education funding councils and the higher education providers to accomplish the statutory requirements set out by the Further and Higher Education Act 1992 and the White Paper: 'Higher Education: A new framework'. The dataset compiles academic data from all higher education institutions including all publicly and privately funded institutions, as well as other organizations that offer higher education courses, including those that are not publicly funded. It includes information about staff members who are employed on an academic contract, such as 'teaching staff only', 'research staff only', and 'teaching and research' staff employed in any type of employment (i.e. part-time or full-time).

The unit of observation in the data is the individual academic staff. The following sample restrictions are imposed throughout the analysis. First, individuals are included if they hold a full-time, permanent teaching and research contract. We consider these as the standard academic contracts, allowing us to have a clearer understanding of wage determination for this type of contract. For example, research-only contracts may be funded by a body outside the higher education provider and may not necessarily follow the academic pay structure. Second, we exclude academics employed in clinical departments as their salary is partly determined by the National Health Service (NHS) in the UK and follow a different pay structure to higher education providers.

Third, we focus on academics in Russell Group institutions.[13] Doing so results in a more homogeneous group of institutions to study the impact of the intervention: the Russell Group includes the UK's research-intensive universities, forming 74 per cent of the research income of all UK institutions and rank high in the World University Rankings. As a result — not least because of the Research Excellence Framework[14] — they are more likely to compete highly in international and national markets for hiring staff as compared to less research-intensive universities (McCormack et al., 2014). Despite this competitive recruitment, however, the high level of autonomy universities have over the hiring of academics and the pay setting, in particular at senior levels (Aghion et al., 2010), may arguably increase the level of gender discrimination (Melly & Puhani, 2013). All these factors additionally make Russell Group institutions comparable to US universities that have typically been the focus of studies on gender pay inequalities, thus allowing us to some extent to draw some comparisons of our results.[15]

Our sample consists of 64,770 observations of 10,770 female faculty and 173,145 observations of 25,205 male faculty, in 24 research-intensive universities over 13 years (2004–2016). We use log real annual earnings as the dependent variable to measure the gender pay gap; as we consider a sample under full-time employment, any differences are attributed to differences in earnings rather than working hours. We use age, age squared and highest educational qualification as our primary independent variables in the wage equation, available in HESA. The data do not include information on marital status, presence/number of children or geographical residence location. Table A1 in the Supplementary Online Appendix lists and defines the variables used in the analysis.

Our research question seeks to estimate the effect of the pay transparency intervention on the gender pay gap. To achieve this, we adopt a difference-in-differences like design in which we compare log real annual earnings of men and women to a sharp change around the gender pay publication by *THE*. Specifically, we estimate the following regression:

$$Y_{itj} = a + \beta_1 D_t + \beta_2 (D_t \times F_i) + X_{it} + \eta_j + \delta_t + \gamma_{jt} + \theta_i + \epsilon_{itj} \tag{1}$$

where Y represents log real earnings using 2016 as the base year for individual i, in university j, in year t. D_t is the pay transparency variable taking the value of one for years after the publication of the first pay report in university j in year t; and the value zero otherwise. Our key regressor is $D_t \times F_i$ which captures the effect of the relative impact on females compared to males following the intervention; a positive coefficient of β_2 indicates a decrease in the gender pay gap.

The vector X accounts for time variant socio-demographic factors that may be correlated with earnings, including age, age squared and highest qualification held. We also control for university fixed effects, η_j, to account for unobservable and time invariant university-specific characteristics that may be correlated with earnings, but not with the pay transparency intervention, such as university rank. For example, due to competition within the Russell Group, some universities are likely to offer higher wages compared to others to attract the best academics. The specification also controls for time fixed effects, δ_t, to account for economic conditions that vary over time which

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may influence wages across sectors. We include university-specific time fixed effects, γ_{jt}, that capture unobservable factors by university-year cells. To account for time-invariant unobserved heterogeneity at individual level, we estimate Equation (1) using a fixed effects model, θ_i. All standard errors are clustered at the individual level.

A compelling aspect of estimating the impact of pay transparency reporting on the gender pay gap relies on two main assumptions. The first assumption is that there were no other policies coinciding with-and/or related to-the pay reporting, causing a change in the gender pay gap. Indeed, in 2007, no other intervention was targeting the gender pay gap. At that time, Athena SWAN was the only policy that was implemented to address issues of inequality by focusing on improving female representation and promotions in Science, Technology, Engineering and Mathematics (STEM) disciplines. We nonetheless test this alternative hypothesis by estimating our baseline model on a sample of academics in Non-STEM departments, who are thus unaffected by Athena SWAN.

The second assumption is that the intervention was not anticipated. This is satisfied since *THE* did not make any official announcements prior to the first report. Arguably, even if universities were informed, say, a few months in advance, salaries in this sector are sticky and cannot be adjusted swiftly (see discussion in Section 2). We empirically investigate the anticipated effects by testing for parallel trends in annual earnings prior to the intervention. We also test whether the intervention is endogenous to the prevailing gender pay gap before the intervention.

To investigate the robustness of our results, we conduct several sensitivity tests. First, we estimate Equation (1) using a balanced sample of academics. Second, we measure the intensity of the intervention by splitting the post-intervention dummy variable and considering two interaction effects instead: one treating the first two years of the intervention (2007 and 2008) denoted as the 'announcement' period, A_t, and another denoted as the 'implementation' period, I_t, from 2009 onwards. To this extent, we estimate the following specification:

$$Y_{itj} = a + \beta_1 A_t + \beta_2 (A_t \times F_i) + \beta_3 I_t + \beta_4 (I_t \times F_i) + X_{it} + \eta_j + \delta_t + \gamma_{jt} + \theta_i + \epsilon_{itj} \quad (2)$$

where variables hold the same definition as above. Here, the treatment variable is split into two: variable A_t takes the value of one in 2007 and 2008 when the first report is published, and zero otherwise; and variable I_t takes the value of one from 2009 onwards, and zero otherwise. Each interacted with the female dummy, F_i.

4 | RESULTS

4.1 | Descriptive statistics

Table 2 presents the descriptive statistics of our sample by gender. On average, male academics earn £7162 (approximately 11 per cent) more compared to their female peers. In total, there are 237,915 observations, 72.7 per cent of which are males. Although there is presently about a 3:1 higher representation of male to female academics, this actually reflects an increasing representation: in the 1970s, only 10 per cent of academics were females (McNabb & Wass, 1997). On average, male academics are more likely to have a doctorate and are by 2 years older compared to their female peers. In terms of ethnicity, male and female academics are generally similar.

Figure 1 plots the average annual trend in the gender pay gap in Russell Group Universities. We plot the raw gender pay gap (left) and the gender pay gap after controlling for individual characteristics (right) separately; both of which produce similar trends. In particular, we observe the

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TABLE 2 Descriptive statistics.

	Female		Male		Difference (Male – Female)	
Variables	**Mean**	**S.D.**	**Mean**	**S.D.**	**Mean**	**S.D.**
Log-Real Earnings	10.89	0.244	11.00	0.268	0.109***	0.003
Ethnicity						
White	0.899	0.302	0.909	0.287	0.0107***	0.0039
Black	0.0047	0.0683	0.0058	0.0762	0.0011	0.0009
Asian	0.0652	0.247	0.0620	0.241	−0.0032	0.0033
Other	0.0314	0.174	0.0227	0.149	−0.0086***	0.0022
Age	44.26	8.937	46.85	9.355	2.596***	0.1109
Highest Qualification						
Doctorate	0.811	0.392	0.880	0.325	0.0694***	0.0047
Postgraduate, equivalent	0.134	0.341	0.0768	0.266	−0.0574***	0.004
First Degree, equivalent	0.04	0.196	0.0287	0.167	−0.0112***	0.0023
Below Undergraduate Level	0.0025	0.0499	0.0025	0.0494	−0.0001	0.0007
Other Qualification	0.012	0.109	0.0115	0.107	−0.0005	0.0013
No Qualification	0.0006	0.0248	0.0004	0.0202	−0.0002	0.0002

Note: HESA dataset between 2004 and 2016. There are 64,770 observations consisted of 10,770 female faculty and 173,145 observations of 25,205 male faculty in 24 universities over 13 years. The sample consists of full-time permanent and research and teaching academics. Log real annual earnings adjusted using 2016 CPI index. Annual earnings are censored at the top and the bottom 1% to prevent extreme outliers affecting mean annual earnings. *** p ⟨ 0.01.

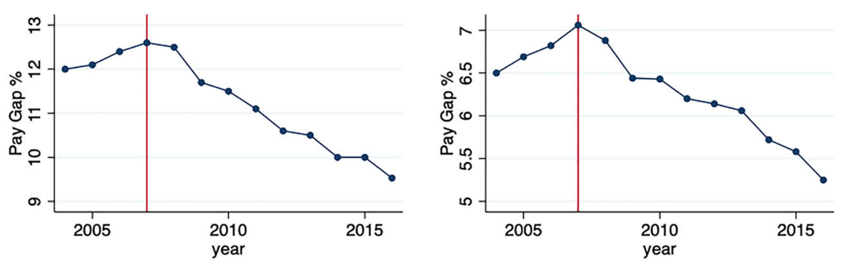

FIGURE 1 Trends in the gender pay gap.
[Colour figure can be viewed at wileyonlinelibrary.com]
Note: Left: Raw data. Right: Adjusted for controls (age, age squared, ethnicity, education and university FE).
Source: HESA 2004–2016.

upward trend up until 2007 — the intervention year — and a gradual decrease thereafter, which is especially pronounced from 2009 onwards. Over the entire sample period, the raw gender pay gap ranged between around 12 per cent and 9.6 per cent, which are comparable to the gender pay gap literature in the UK (Blackaby et al., 2005; Euwals & Ward, 2005; Gamage et al., 2020; McNabb & Wass, 1997; Mumford & Sechel, 2020; Schulze, 2015), the United States (Ginther, 2001; Ginther & Hayes, 2003; Ginther & Kahn, 2014; Sutanto et al., 2014; Tao, 2018) and Europe (Corsi et al., 2014; Hospido et al., 2022).

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TABLE 3 Impact of pay transparency on earnings.

	(1)	(2)	(3)	(4)
	Full sample		Balanced sample	
Dependent Variable: Log-Earnings	**Model 1**	**Model 2**	**Model 3**	**Model 4**
Pay Transparency	0.0303	−0.0172	0.245	0.147
	(0.115)	(0.114)	(0.164)	(0.137)
Pay Transparency × Female	0.0065***	0.0062***	0.0144***	0.0127***
	(0.0019)	(0.0018)	(0.0036)	(0.0033)
Age	0.0776***	0.0798***	0.0548***	0.0619***
	(0.0096)	(0.0095)	(0.0138)	(0.0115)
Age^2	−0.0006***	−0.0006***	−0.0005***	−0.0006***
	(0.0001)	(0.0001)	(0.0001)	(0.0001)
Postgraduate	0.0032	0.0022	−0.006	−0.0007
	(0.0033)	(0.0032)	(0.0077)	(0.0071)
First Degree	−0.012**	−0.0077	−0.0211	−0.0143
	(0.0054)	(0.0047)	(0.0126)	(0.01)
Below Undergraduate Level	0.0059	0.0071	0.065	0.0504
	(0.0225)	(0.0177)	(0.0551)	(0.0395)
Other Qualification	−0.0157**	−0.0192***	−0.0113	−0.0199**
	(0.0062)	(0.006)	(0.011)	(0.0099)
Constant	8.677***	8.632***	9.541***	9.303***
	(0.382)	(0.377)	(0.57)	(0.474)
University FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Uni × Year FE	No	Yes	No	Yes
N	237,915	237,915	67,910	67,910
N of Individuals	35,975	35,975	5225	5225
R^2	0.568	0.628	0.586	0.667

Note: The table presents estimates from Equation (1). Dependent variable is log real annual earnings adjusted using 2016 CPI index. Annual earnings are censored at the top and the bottom 1% to prevent extreme outliers affecting mean annual earnings. Robust standard errors clustered at the individual level are reported in parentheses. Columns 2 and 4, offering a full set of controls, is our benchmark specification. *** $p < 0.01$; ** $p < 0.05$.

4.2 | Main results

Table 3, columns 1 and 2, presents the estimates of Equation (1) using a fixed-effect model. The first column presents a specification without controlling for university-specific time fixed effects, while column 2 presents the full specification (our benchmark specification). We estimate several other specifications by adding controls progressively to confirm that the interaction variable is consistent throughout these specifications (results available upon request). According to this benchmark estimate, following the pay reporting, the log of female annual earnings increased, on average, by 0.62 percentage points, while the log wage of males remained constant; this corresponds to an increase in average female annual earnings by £323 relative to men. Based on average pre-intervention earnings levels of men and women, this translates to a 4.38 per cent

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decrease in the gender pay gap.[16] We do not find a significant change in the overall earning levels of academic staff.

The remaining coefficients in Table 3 are as expected: on average, an increase in age increases log-annual earnings by 8 percentage points, diminishing as an individual gets older; academic staff with doctorates and postgraduate qualifications earn significantly more compared to academic staff with other levels of qualification.[17] As racial background is a fixed effect, we estimate instead interaction terms between a non-White indicator (i.e. Asian, Black and other background) with the female (F_i) and pay transparency (D_t) variables, as well as with our key regressor, $D_t \times F_i$. Table A2 in the Supplementary Online Appendix suggests that the ethnicity interaction terms do not have a significant impact on log-earnings.

Table 3, columns 3 and 4, estimates Equation (1) using a balanced sample to assess the sensitivity of our findings. Despite the expected attrition in our data due to the 13-year span and the focus on Russell Group institutions, the results of the balanced sample are in line with the previous findings. In fact, the estimates are now more pronounced, with log female annual earnings estimated to increase by 1.27 percentage points, while male earnings remain constant. This is a fall in the gender pay gap of 11.43 per cent based on average male and female wage gap before the intervention.[18] We also do not find a significant change in the overall earning levels.

Table 4, column 1, presents the estimation results of Equation (2) (for the model including all controls). We only report the coefficients of the intervention terms and their interaction with the female dummy variable. The announcement–female interaction is not statistically significant, whereas the implementation–female interaction is positive and highly significant. Following 2009, log female annual earnings increase by 0.76 percentage points compared to male annual earnings, which is 0.14 percentage points (18.4 per cent) higher than the estimate of our benchmark specification (i.e. 0.62 percentage points in Table 3, column 2). This confirms our hypothesis that wages take time to adjust following implementation and, after accounting for the adjustment period, our results are similar but larger in magnitude compared to those of the benchmark specification.

4.2.1 | Comparison to previous causal estimates

To complement our study, we compare our results to research findings that examine a causal effect of pay transparency intervention on the gender pay gap. Table 1 presents a summary of the findings of other relevant studies. Existing studies (four out of five) find a fall in the gender pay gap following a pay transparency intervention, which is similar to our results.[19] However, the magnitudes reported in these are larger than ours: whereas, previous relevant studies find a fall in the pay gap between 2 and 2.8 percentage points, we see a fall in the gender pay gap by 0.62 percentage points. For example, papers studying the causal effect of a 2017 pay transparency intervention in the UK that affected all firms with more than 250 employees find their effects in the magnitude of 2.3 and 2.8 (Duchini et al., 2022a).

A plausible explanation for our smaller estimate may be that the intervention investigated here only affected the higher education sector; consequently, we observe lower wage gap outcomes (after controlling for occupational rank) than studies that include all industries and occupations. Another justification may be that wages are more difficult to adjust in the UK higher education sector. As highlighted in Section 2.1, pay scales of academics below professorial rank — which forms a relatively large proportion of the academic population — are governed by a sector-wide collective bargaining agreement. Our evidence from senior versus non-senior academics suggests

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TABLE 4 Further specifications.

Dependent Variable: Log-Earnings	(1) Intensity of Intervention	(2) Non-STEM Sample
Announcement	0.0487	
	(0.038)	
Announcement × Female	−0.0024	
	(0.0014)	
Implementation	−0.0175	
	(0.114)	
Implementation × Female	0.0076***	
	(0.0022)	
Pay Transparency		0.22**
		(0.0732)
Pay Transparency × Female		0.0068**
		(0.0032)
Constant	8.634***	9.274***
	(0.377)	(0.238)
Individual Controls	Yes	Yes
University FE	Yes	Yes
Year FE	Yes	Yes
Uni × Year FE	Yes	Yes
N	237,915	78,575
N of Individuals	35,975	12,755
R^2	0.628	0.629

Note: Column 1 presents estimates based on Equation (2). Column 2 presents estimates based on Equation (1) for selected non-STEM disciplines, including: Business and Management, Social Sciences (law, economics, social work & social policy, sociology, politics & international studies, anthropology & development studies) and Humanities and language-based studies (area studies, theology & religious studies, philosophy, classics, history, English language & literature). The dependent variable is log real annual earnings adjusted using 2016 CPI index. Annual earnings are censored at the top and the bottom 1% to prevent extreme outliers affecting mean annual earnings. Individual controls include age, age^2 and highest level of education. Robust standard errors clustered at the individual level are reported in parentheses. *** $p < 0.01$; ** $p < 0.05$.

that this may be a factor. In fact, when we consider those at professorial level who are not subject to pay structures, then the estimates are closer to what other studies find for the UK (see Section 5.2). Finally, the lower effect we find may be further reinforced by the fact that there are low levels of awareness among academics of this publicly available data (Pfefer, 2020).

4.3 | Identification

The key identification assumptions underlying our approach is that this pay transparency intervention was uncorrelated with pre-existing differences in the earnings of men and women in this sector (after controlling for time-varying controls, time-invariant university characteristics, a common trend and university-specific time trends) and that our findings are not driven by other sharp

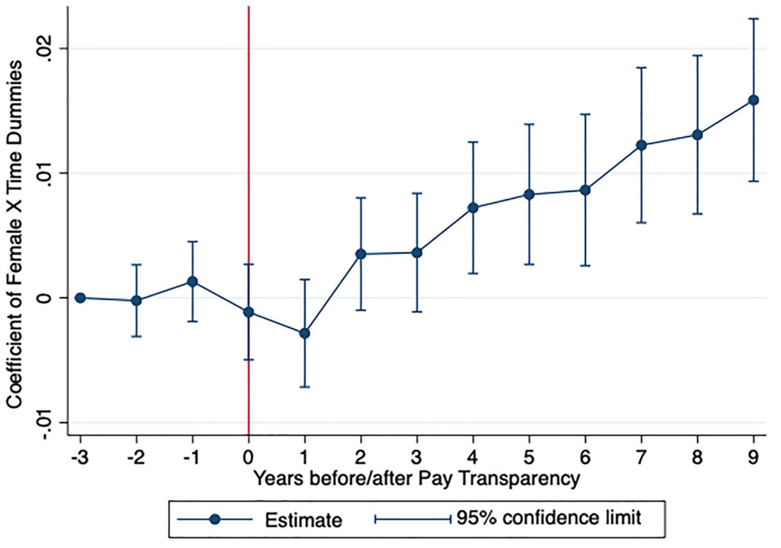

FIGURE 2 Coefficient plot of estimates from Equation (3).
[Colour figure can be viewed at wileyonlinelibrary.com]

Note: Sample: Full-time permanent academics on teaching and research contracts observed over 13 years between 2004 and 2016. The figure presents coefficient estimates of Female and year dummy interaction in Equation (3). The dependent variable is log real annual earnings adjusted using 2016 CPI index. Annual earnings are censored at the top and the bottom 1 per cent to prevent extreme outliers affecting mean annual earnings.
Source: HESA dataset.

changes around the time of the intervention. To assess the plausibility of these assumptions, we proceed in several ways.

First, we exploit the exact timing of intervention to test whether the pay transparency shock generated sharp changes in the gender pay gap, which are arguably orthogonal to unobserved determinants of the wage gap that evolve more smoothly. Following Autor (2003), Kavetsos et al. (2021) and Kleven et al. (2019), we employ a difference-in-differences event study methodology, where we substitute the intervention variable, D_t, with a complete set of dummies going 3 years before to 9 years after the pay publication. We interact each of these time dummies with the female dummy. In particular, we estimate:

$$Y_{itj} = a + \sum_{\tau=-3}^{\tau=9} \theta_\tau D_\tau + \sum_{\tau=-3}^{\tau=9} \beta_\tau (D_\tau \times F_i) + X_{it} + \theta_i + \eta_j + \gamma_{jt} + \epsilon_{itj} \tag{3}$$

where D_τ is a vector of year dummies for the τ_{th} year before and after the pay reporting, capturing the general trends common to both men and women before and after the intervention. To separate the gender differences in the trends, we interact the time dummies with the female dummy, $D_\tau \times F_i$. We also include the same set of controls and fixed effects given in Equation (1). We use the first observation in our sample, year(-3), as the reference year.

Figure 2 plots the estimated coefficient of the interaction term (Female×Time dummies) together with the 95 per cent confidence bands. Table A3 in the Supplementary Online Appendix presents the estimated values.[20] Figure 2 shows that the decline in the gender pay gap does not occur before the pay transparency intervention: none of the coefficients for the years preceding the pay publication are statistically significantly different from zero.[21] This result

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increases our confidence in the validity of our identification strategy, as it would be difficult to explain the discontinuous fall in the gender pay gap following pay publication due to trends in unobservable characteristics.

Second, we use a sample of non-STEM departments to eliminate the possibility that our results are driven by a policy that might have coincided with pay publication. One such policy is the Athena SWAN intervention. The Athena SWAN Charter was established by the Equality Challenge Unit (ECU) in 2005, accrediting universities and departments with Bronze, Silver or Gold status for their commitment to advancing gender equality in the STEM disciplines. The charter does not specify targets for universities/departments. Instead, it requires universities to conduct an in-depth assessment of gender equality within the university/department and implement an action plan to improve gender equality.

Before 2015, the charter focused on increasing female representation and promotion, not the gender pay gap. This means it is, in principle, not related to wages. Yet, to net out the effect of the Athena SWAN empirically, we follow Gamage and Sevilla (2019) by estimating the impact of the pay transparency initiative on a non-STEM sample for which the Athena SWAN intervention does not hold, since the Athena SWAN was extended to non-STEM departments in May 2015. We estimate the baseline model given in Equation (1) on a comparable non-STEM departments within our sample, including Business and Management, Social Sciences (including law, economics, social work & social policy, sociology, politics & international studies, anthropology & development studies) and Humanities and Language-based studies (including area studies, theology & religious studies, philosophy, classics, history, English language & literature).[22]

The results presented in Table 4, column 2, show that the estimates are similar to the baseline results. On average, following the pay reporting, log female annual earnings increased by 0.68 percentage points, equivalent to the 0.62 percentage points when considering the full sample in Table 3. This result confirms that the pay transparency intervention affected the gender pay gap in non-STEM disciplines that are not affected by the Athena SWAN intervention. It, thus, provides compelling evidence suggesting that our results are not driven by another policy that coincided with pay publication.[23]

5 | MECHANISMS BEHIND THE PAY TRANSPARENCY

Why would the gender pay gap fall? We identify two main channels noted in the literature, the reputation effect and the bargaining effect, which we consider in detail.

5.1 | Reputation effect

Our main specification exploits the temporal variation of the policy intervention by comparing the impact on women compared to men. Here, we test whether the impact on the differential effect on earnings is influenced by the intensity of the gender pay gap before the intervention. 'Elite' institutions, such as the Russell Group universities, are aware of their standing in the sector and are arguably concerned in maintaining that image. These organizations could, for example, be more vigilant about potential reactions associated with the wages they offer, which could damage their reputation (Graffin et al., 2013; McDonnell et al., 2015; Rhee & Haunschild, 2006). It is, thus, interesting to investigate empirically whether universities with high gender pay gaps pre-intervention adjust female pay to become more similar to universities that have a lower pay gap.

TABLE 5 Test for the reputation effect.

Dependent Variable: Log-Earnings	Pre-existing pay gap
Pay Transparency	−0.0266
	(0.114)
Uni Gap	−0.0905
	(0.0533)
Pay Transparency × Female	0.0062***
	(0.0018)
Pay Transparency × Uni Gap	−0.016
	(0.0222)
Female × Uni Gap	0.0072
	(0.0087)
Pay Transparency × Female × Uni Gap	−0.0005
	(0.0016)
Constant	8.576***
	(0.374)
Individual Controls	Yes
University FE	Yes
Year FE	Yes
Uni × Year FE	Yes
N	237,915
N of Individuals	35,975
R^2	0.628

Note: The dependent variable is log real annual earnings adjusted using 2016 CPI index. Table presents estimates based on Equation (1) after interacting average standardized university level pre-existing gender pay gap with the female and intervention interaction term. Individual controls include age, age^2 and the highest level of education. Robust standard errors clustered at the individual level are reported in parentheses. $^{***}p < 0.01$.

To this end, our identification strategy exploits the heterogeneity in the pre-intervention gender pay gap at the university level to study whether universities with a high pre-intervention wage gap react more aggressively than universities with a low pre-intervention wage gap.[24] To do so, we introduce the university-level pre-intervention wage gap (Uni Gap) to Equation (1) and form a triple interaction between the university-level pre-intervention wage gap, the intervention (D_t) and female dummy (F_i). We measure the university-level pre-intervention wage gap by taking the difference between the male and female log annual earnings at the university-year level, averaged over the pre-treatment period. We standardize the pre-intervention earnings gap to have a mean zero and a standard deviation of one.

Table 5 presents the results. These suggest there are no differences of the pay transparency initiative across universities depending on a university's pre-intervention pay gap. We, therefore, do not have sufficient evidence in support of a reputation effect. Repeating, however, this estimation for the five leading UK universities in our sample ('Top 5'), based on the QS World University Rankings (i.e. U of Cambridge, U of Oxford, Imperial College London, University College London and U of Edinburgh), supports the hypothesis of a reputation effect; see Table A5 in the Supplementary Online Appendix. Here, the transparency initiative led to the adjustment of all salaries, especially those of females, with further positive and statistically significant

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adjustments given their pre-intervention wage gap (with both Pay Transparency×Uni Gap and Pay Transparency×Female×Uni Gap being positive and statistically significant).

5.2 | Bargaining effect

The literature suggests that when presented with wage information, women bargain better. For example, using data from an online recruitment platform, Roussille (2022) finds that while women ask for a much lower initial salary compared to men, they raise their bid when informed about the median salary for their position. We conduct two tests to investigate whether our results are driven by women bargaining for higher wages.

First, we explore whether the effect of the policy differed by universities depending on the female composition pre-intervention. Presented with the wage information, universities with high female representation may face greater pressure to adjust wages compared to universities with low female representation; in which case we might witness a greater fall in the gender pay gap. To test this hypothesis, we exploit the heterogeneity in the pre-intervention female representation at the university level, averaged over the pre-treatment period. We standardize this variable to have a mean of zero and a standard deviation of one, which we interact with the $D_t \times F_i$ variable given in Equation (1). Table 6, column 1, presents the results. The coefficient of the triple interaction is not statistically significant. This indicates no significant difference in the outcome of universities with high female representation, suggesting that there is no bargaining effect taking place at the university level.

This result is not driven by universities with more women paying women less and having less money to give out to them. Our sample of Russell Group institutions arguably define ‘the elite’ in the UK academic sector: they attract about 3/4 of the research income across all UK academic institutions and compete highly, as compared to other non-research-intensive institutions, to attract ‘research and teaching’ staff. Hence, Russell Group institutions being resourced-constrained to give out to lower-paid female staff is not a concern that might be driving this result.

Second, we exploit the pay structure of the UK higher education sector to test whether the impact on the wage gap is different across the wage distribution. Section 2.1 details the differences in wage determination between professorial and non-professorial ranks. A sector-wide collective bargaining agreement determines the pay of academics below the professorial rank, with a wage floor/ceiling bound for each rank. On the other hand, professorial pay is not regulated, offering greater room for professors to bargain a higher wage in response to the information shock than academics below professorial rank.

The data, however, do not allow us to identify academic ranks before 2009 (i.e. before the introduction of the intervention). Therefore, we use instead earnings above and below the median as a proxy for occupational hierarchy.[25] We estimate Equation (1) on academics below the median earnings and those above the median earnings separately. The results in Table 6, columns 2 and 3, show that on average female academics earning below the median value (column 2) experience an increase in log earnings by 0.50 percentage points (a nominal fall of £223 or a 22.8 per cent fall in the gender pay gap).[26] More importantly, academics earning above the median annual earnings (column 3) experience a 1.43 percentage point fall in gender pay gap (a nominal fall of £980 or a 25.9 per cent fall in the earnings gap), which is almost three times the fall experienced by academics earning below the median annual earnings.[27] The difference in the impact between academic staff below and above the median is statistically significant at the 95 per cent

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TABLE 6 Test for the bargaining effect.

Dependent Variable: Log-Earnings	(1) Female repres.	(2) Below median wage	(3) Above median wage	(4) Movement
Pay Transparency	−0.0131	0.237***	0.0434	0.126
	(0.114)	(0.0709)	(0.011)	(0.0805)
Pay Transparency × Female	0.0063***	0.005**	0.0143***	0.0011
	(0.0018)	(0.0021)	(0.0024)	(0.0029)
Avg Rep	0.04***			
	(0.0149)			
Pay Transparency × Avg Rep	−0.0055			
	(0.0085)			
Female × Avg Rep	−0.0135			
	(0.0083)			
Pay Transparency × Female × Avg Rep	−0.002			
	(0.0019)			
Constant	8.61***	9.209***	9.449***	0.505
	(0.378)	(0.204)	(0.411)	(0.268)
Individual Controls	Yes	Yes	Yes	Yes
University FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Uni × Year FE	Yes	Yes	Yes	Yes
N	237,915	111,850	113,005	237,915
N of Individuals	35,975	23,620	18,535	35,975
R^2	0.628	0.719	0.567	0.325

Note: The dependent variable is log real annual earnings adjusted using 2016 CPI index. Column 1 presents estimates based on Equation (1) after interacting average standardized university-level female representation (Avg Rep) with the female and intervention interaction term. Columns 2 and 3 present estimates for Equation (1) for academics below and above the median wage, respectively. Column 4 presents estimates based on Equation (4), where the dependent variable is binary, taking value 1 if an individual moves university or leaves academia. Annual earnings are censored at the top and the bottom 1% to prevent extreme outliers affecting mean annual earnings. Individual controls include age, age^2 and the highest level of education. Robust standard errors clustered at the individual level are reported in parentheses. *** $p < 0.01$; ** $p < 0.05$.

significance level. These estimates suggest women primarily drive the effect at the top of the earnings distribution. The results of a triple interaction for those above median wage offer consistent results (see Table A6 in the Supplementary Online Appendix).

5.2.1 | Movement

The results on bargaining suggest that negotiation effects drive the reduction in the gender pay gap results. These negotiations arguably occur in two ways: (a) within institutions, where females negotiate higher salaries within the rank and through promotions; or (b) females negotiate a higher wage through outside offers. To investigate this, we exploit the panel nature of the administrative data to unravel whether the fall in the gender pay gap results from an increase in departure

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rates of female academics. To this extent, we examine whether female academics are more likely to move between universities or out of the sector, or whether they stay and progress within their institution following the pay publication. We estimate the following linear probability model:

$$M_{itj} = a + \beta_1 D_t + \beta_2(D_t \times F_i) + X_{it} + \theta_i + \eta_j + \delta_t + \gamma_{jt} + \epsilon_{itj} \tag{4}$$

where M represents movement, defined as a dummy variable equal to one if the individual moves to a different university or leaves the higher education sector in year t+1, and zero otherwise. D_t is the pay transparency variable taking the value of one after the publication of the first pay report in university j in year t, and the value of zero otherwise. Our key regressor is $D_t \times F_i$ which captures the effect of the relative impact on females compared to males following the intervention; a positive coefficient of β_2 indicates that women are more likely to move relative to men.

Table 6, column 4, presents these results, which should be viewed with a note of caution because of two reasons. First, although we find that following the pay transparency intervention, women are more likely to move to other universities or outside academia compared to men, this estimate is statistically insignificant. Second, given the focus on academics in Russell Group institutions, we are not able to distinguish between movement to non-Russell Group institutions and exit from the sector. Overall, we provide suggestive evidence of a fall in the pay gap being driven by women negotiating a higher salary within the same university.

6 | CONCLUSION

We study the impact of a pay transparency intervention on the gender pay gap in the UK higher education sector. We find that following the introduction of the pay transparency, the gender pay gap narrows significantly. An increase in female earnings drives this decrease in the gap. Further estimations allow us to rule out the possibility of other factors driving these results. We also demonstrate that senior academics in the upper half of the earnings distribution mostly drive this decrease in the earnings gap. In general, we do not find pre-existing gender pay gaps or the gender composition at university level to influence the impact on gender pay inequality; the exception being when we consider the leading UK Universities (‘Top 5’), supporting a reputation effect. Thus, the overall evidence suggests that bargaining effects tend, more generally, to drive our results. Further exploration of the panel nature of our data reveals that the bargaining effects likely take place within the given institution.

We are not able to examine how the intervention affects academics in specific ranks within the non-full-professorial staff (i.e. Lecturer, Senior Lecturer and Reader; and equivalently, Assistant Professor and Associate Professor for institutions that have adopted the US academic convention) due to academic rank markers missing from the data. Although the evidence suggests a fall in the earnings gap following the intervention, we cannot rule out any spill-over effects that may limit the impact of pay transparency policy on the gender pay gap (whereby men take advantage of pay information to increase their earnings).[28] To this extent, our results can be interpreted as lower bound estimates of the true effect of the pay transparency intervention in the academic sector.

In addition, the literature has raised concerns about the awareness and accessibility of such transparency policies: if individuals are unaware of policies, or if accessibility is difficult and/or costly, the impact on wage levels and, in particular, on gender pay gaps will be sub-optimal (Baker et al., 2023). Data on academics' union membership and the gender of the negotiation counterparty could potentially offer some insight around this issue; in the absence of such information in

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our data, we are, however, unable to undertake any further analysis towards this direction, which could be the focus of future research.

The gender pay gap is a prominent topic of discussion among policymakers and academics. Various governments use pay transparency policies as an instrument to address gender pay inequality. Our study adds to this literature, suggesting that such a low-cost intervention may lead to considerable reductions in the gender pay gap.

ACKNOWLEDGEMENTS

The authors wish to thank Ian Gregory-Smith, Julian Jessop, Christian Krekel, Ilias Kyriopoulos and participants at the Centre for Globalisation Research 2020 seminar series, the Scottish Economic Association 2021 Conference, the Bocconi 2021 Conference and the 2021 Conference on Diversity and Inclusion in Economics, Finance and Central Banking, jointly organized by the Bank of Canada, Bank of England, European Central Bank and the Federal Reserve for valuable comments and feedback in earlier drafts of this study. The Editor and two anonymous reviewers provided excellent comments and suggestions.

DATA AVAILABILITY STATEMENT

The data that support the findings of this study are collected from the Higher Education statistical Agency (HESA), available in the public domain: www.timeshighereducation.com/features/times-higher-education-pay-survey-2015/2019360.article.

ETHICS STATEMENT

Not applicable as the research is based on secondary data.

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ENDNOTES

[1] Fawcett Society launched the ‘Right to Know’ campaign to push for a new equal pay bill to modernize the law on equal pay. They find that 53 per cent of women and 47 per cent of men at work are reluctant to share information with colleagues on how much they earn. Further, 31 per cent of employees believe that the contracts prohibit them from talking about their pay. See https://www.fawcettsociety.org.uk/right-to-knowhttps://www.fawcettsociety.org.uk/right-to-know.

[2] Refer: https://women.govt.nz/documents/country-case-studies-pay-transparency.

[3] Related, see Azmat et al. (2019) for a grade transparency intervention at a Spanish higher education institution.

[4] This can arguably be linked to evidence suggesting that women develop differences in their willingness to compete that emerge quite early in life and persist (Sutter & Glätzle-Rützler, 2014), which can, however, be minimized if this gender difference in the willingness to compete is communicated via an information intervention (Kessel et al., 2021).

[5] *THE* publication takes place in April of every calendar year and is publicly accessible through the *THE* website. See https://www.timeshighereducation.com/featureshttps://www.timeshighereducation.com/features.

[6] *THE* had to formally obtain permission to publish the pay data to comply with data protection regulations at the time, and informed universities a couple of months before the release of the publication.

[7] See also Chari and Goldsmith-Pinkham (2017) for gender representation in economics conferences, and Apesteguia et al. (2011) and Baltrunaite et al. (2014) for experimental evidence focusing on inequalities in gender representation.

[8] Several other studies focus on the impact of pay transparency on other factors such as employee well-being (Perez-Truglia, 2020), job satisfaction (Akerlof & Yellen, 1990; Breza et al., 2018; Card et al., 2012), work effort, output and employee relation (Cullen & Perez-Truglia, 2022).

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[9] See https://www.ucu.org.uk/frameworkhttps://www.ucu.org.uk/framework.

[10] Some universities might have spine points for Professors but, as this practice is unregulated and is university-specific, such data are not included in the Higher Education Statistical Agency (HESA) where our data are obtained from.

[11] The THE obtained the data for the reports from the Higher Education Statistical Agency data.

[12] See 2015 report https://www.timeshighereducation.com/features/times-higher-education-pay-survey-2015/2019360.article https://www.timeshighereducation.com/features/times-higher-education-pay-survey-2015/2019360.article.

[13] These are: University of Birmingham, University of Exeter, University of Bristol, University of Liverpool, University of Cambridge, King's College London, Cardiff University, University of Nottingham, Durham University, University of Southampton, University of Edinburgh, University of York, University of Glasgow, Imperial College London, University of Leeds, London School of Economics, University of Manchester, Newcastle University, University of Oxford, Queen Mary University of London, University of Sheffield, Queen's University Belfast, University College London and University of Warwick.

[14] Research Excellence Framework in the UK is a system of assessing the quality of research output. It focuses on the following three dimensions: the quality of outputs, their impact beyond academia and the environment that supports research. For more information, see https://www.ref.ac.uk/https://www.ref.ac.uk/.

[15] We considered examining the effect of the intervention on other UK higher education institutions. Although the estimated results support the positive effect the intervention had in reducing the gender pay gap, there appears to be substantial institutional heterogeneity preventing us to uncover a common trend between male and female earnings prior to the intervention.

[16] Pre-treatment average female annual earning is £52,274 and average male annual earning is £59,646.

[17] As observed in Table 2, those with only a UG degree, no qualification or 'other qualification' are a small proportion of the sample. Estimating Equation (1) by omitting individuals in these categories does not change the results, with the coefficient for Pay Transparency×Female estimated at 0.0064 (from 0.0065 in Table 3, column 2) and 0.013 (from 0.0127 in Table 3, column 4), respectively; both significant at the 1 per cent level.

[18] Pre-treatment average annual female earning is £51,520 and average annual male earning is £57,161

[19] Table 1 also includes literature on pay disclosure bans and its impact on gender pay gap. However, we consider this type of intervention to be different to pay transparency albeit they improve the information asymmetry in the labour market, therefore, we have not compared our results to these studies.

[20] Note that the delay in statistically significant increases in female earnings is consistent with the evidence presented in Table 4.

[21] Switching the reference year — for example to that of the year before the intervention — does not change Figure 2; rather, re-scales it.

[22] A balance test reveals that, on average, the pay gap and the earning are not statistically different (see Table A4 in the Supplementary Online Appendix).

[23] Note that we do not consider gender equity plans required by some funding bodies in order to award the funds, such as 'Horizon 2020' (the predecessor of 'Horizon Europe'), to be a threat to our identification approach. First, 'Horizon 2020' ran between 2014 and 2020; our sample ends in 2016. Second, the occurrence/timing of Horizon funding varies by institution: it may be Departmental/School-specific and may thus not be transparent across the whole University, let alone the wider sector; there is time variation in applications by Departments/Schools and thus time variation in the equity plan put in place.

[24] This identification strategy follows Card (1992), who studies the impact of federal minimum wage by exploiting the variation in the state-specific minimum wage level before the roll-out of the federal minimum wage. Bertrand et al. (2019) and Maida and Weber (2020) use a similar identification strategy to investigate the impact of gender quota laws on corporate boards.

[25] To verify that professors are captured above the median earnings, we look at professorial starting annual pay for the period when we do have 'professorial markers', by using the pay information published by the universities. For further validation, we surveyed a sample of 14 professors within our network. This exercise suggests that starting professorial earnings are close to the median value of the earnings distribution.

[26] Pre-intervention average gender pay gap below the median earnings is 2.16 per cent (equivalent to £979).

[27] Pre-intervention average gender pay gap above the median earnings is 5.23 per cent (equivalent to £3783).

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[28]The estimate of 'pay transparency' is only positive significant when we consider non-STEM Departments (Table 4) and below-median earnings (Table 6) and the 'top 5' institutions (Table A5 in the Supplementary Online Appendix), which can be interpreted as such.

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SUPPORTING INFORMATION

Additional supporting information can be found online in the Supporting Information section at the end of this article.

How to cite this article: Gamage, D.K., Kavetsos, G., Mallick, S. & Sevilla, A. (2023) Pay transparency intervention and the gender pay gap: Evidence from research-intensive universities in the UK. *British Journal of Industrial Relations*, 1–26. https://doi.org/10.1111/bjir.12778

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