Droughts beyond the farm: Gendered labour market responses and sectoral spillovers in a diversified economy

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Abstract

Droughts are becoming more frequent under climate change, yet most evidence on their labour market consequences comes from low income settings dominated by subsistence agriculture. How diversified economies absorb these shocks, and how the burden distributes by gender, is far less understood. I use quarterly microdata from Chile's National Employment Survey (2010–2024) matched to drought exposure at region level from the Standardised Precipitation Evapotranspiration Index to estimate a stacked event study of effective hours, sectoral employment shares, and contract composition by gender. Three findings stand out. First, adjustment within agriculture operates through the extensive margin for women. Their agricultural employment share falls by 0.8 percentage points on impact and 1.4 one quarter later, while men's share and continuing workers' hours barely move, consistent with non-renewal of seasonal contracts. Second, shocks propagate beyond agriculture. In regions with above median agricultural dependence, services hours decline by 2–5 per week and remain depressed for at least ten quarters, with sharper medium horizon losses in manufacturing. Third, drought raises interregional out migration, attenuated in more diversified origin regions, yet women who stay see no earnings gains. These findings establish that, in diversified labour markets, climate shocks widen gender gaps mainly through participation, job quality, and local non-agricultural spillovers, margins that wage and hours analyses miss.

Keywords: drought, labour markets, gender, agriculture, migration, Chile

JEL codes: J16, J43, O13, Q54, R23

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1 Introduction

Droughts are becoming more frequent and severe as the climate warms, with direct implications for the agricultural sector as they act as negative productivity shocks that reduce yields, farm revenues, and threaten rural livelihoods (Mueller and Osgood, 2009; Dell et al., 2014; Jayachandran, 2006; Burgess et al., 2017; Afridi et al., 2022; World Bank, 2016). In economies where agriculture is not the dominant employer but remains tightly integrated with manufacturing and services, these effects can reshape labour market outcomes far beyond the farm sector. Lower agricultural output and income can contract labour demand in agriculture and propagate to linked activities such as processing, transport, retail, and other local services (Mueller and Quisumbing, 2011; Emerick, 2018; Bastos et al., 2013; Branco and Féres, 2021). These consequences may also unfold differently for women and men. Given that they are often concentrated in different tasks and contractual arrangements, and face different constraints on mobility and time allocation, the same shock can generate sharply different trajectories (Mahajan, 2017; Afridi et al., 2022; Eastin, 2018; Food and Agriculture Organization, 2023). As droughts intensify under climate change, it has become increasingly important to understand how agricultural shocks affect labour markets in diversified economies, where sectoral linkages, local economic structure, and gender may shape both direct and indirect effects.

This paper studies these dynamics through three related research questions. First, within agriculture, I examine whether severe drought onsets alter hours among employed workers, sectoral employment shares, and the stability of employment contracts, and whether these responses differ by gender. Second, I study whether drought shocks generate spillovers to manufacturing and services in regions where local employment is more dependent on agriculture, isolating differential responses through baseline agricultural dependence which captures local demand and production linkages. Third, I examine whether internal migration responds to drought onset and whether regions with more diversified local economies experience smaller spatial adjustment, suggesting that diversification buffers workers against the need to relocate. The distinction between margins based on contracts (non-renewal of seasonal matches, shifts from open ended to limited term arrangements) and standard wage or hours adjustment is central to the analysis. In seasonal agricultural labour markets, adjustment need not operate through proportional hour cuts or wage declines; instead, it can manifest through selection out of employment and changes in contract composition, margins that are rarely observed in existing work but that are first order determinants of job quality and income stability.

A growing body of work shows that drought impacts are not gender neutral within

agriculture, because women and men are segregated into different tasks and jobs, and face different constraints on mobility and time allocation (Mahajan, 2017; Afridi et al., 2022; Eastin, 2018; Van Aelst and Holvoet, 2016; Food and Agriculture Organization, 2023). Evidence from agrarian settings highlights that women often experience sharper declines in paid employment and weaker recovery, while men are more able to relocate across activities or migrate. However, two gaps remain for more diversified labour markets. First, less is known about whether gender differences operate through margins based on contracts, such as non-renewal of seasonal matches or shifts from open ended to limited term arrangements, rather than through contemporaneous wage or hours adjustments. Second, there is limited evidence on how gendered adjustment within agriculture interacts with spillovers to non-agricultural sectors that employ many women and that are connected to farm income and volumes (Barrientos and Visser, 2012; Bastos et al., 2013; Branco and Féres, 2021).

A complementary strand emphasizes mechanisms operating beyond agriculture. Weather shocks can propagate through local general equilibrium effects, including contractions in demand for locally produced services and disruptions to processing, storage, and transport activities linked to farm output (Bastos et al., 2013; Branco and Féres, 2021; Aragón et al., 2021). Much of this work focuses on aggregate employment and wages, leaving limited evidence on how spillovers vary by gender and by the contract structure of affected jobs. It is also less clear whether non-agricultural responses operate mainly through local demand effects or through production linkages across sectors, making it difficult to assess which types of workers and activities are most exposed when agricultural shocks propagate through the local economy.

A third strand studies spatial adjustment, focusing primarily on migration and mobility responses to climate shocks, including both internal and international movements. Existing evidence shows that adverse climate shocks can increase migration by reducing agricultural income and local employment opportunities, although these responses are selective, costly, and shaped by household resources, migration networks, and conditions in origin and destination labour markets (Hornbeck, 2012; Feng et al., 2010; Baez et al., 2017b,a; Nawrotzki et al., 2017). This evidence highlights migration as an important adjustment margin, but says less about when it substitutes for, or interacts with, other forms of labour market adjustment. In economies with stronger institutions and a more diverse sectoral base, migration may be a less common response because workers can adjust across sectors or jobs within the same local economy. However, less is known about how local economic structure mediates the labour market effects of drought, or whether out migration stabilizes wages for those who remain when drought depresses local demand more broadly.

These three bodies of work have developed largely in parallel, in part because the mech-

anisms linking weather variation to farm level productivity differ from those linking gender segmentation to contract outcomes, and both differ from those governing migration decisions. However, the central question lies precisely at their intersection. When drought hits a diversified economy, workers may adjust through direct displacement within agriculture, indirect contractions in linked non-agricultural sectors, and spatial reallocation, with each of these channels shaped by gender. Existing evidence still characterizes adjustment mainly through wages, hours, and employment, offering limited guidance on job quality margins central to seasonal labour markets, on how gendered adjustment within agriculture interacts with spillovers to non-agricultural sectors that employ many women, and on whether local economic structure mediates the need for spatial adjustment. The literature leaves open how these margins combine to shape the distributional effects of drought in diversified labour markets.

To answer the research questions, I link quarterly individual level microdata from Chile's National Employment Survey (ENE) for 2010–2024 to region level drought measures based on the three month Standardised Precipitation Evapotranspiration Index (SPEI-3). Chile offers an informative setting for three reasons. First, the country has experienced a well documented megadrought since 2010, with substantial heterogeneity across regions and quarters, and rich survey microdata at worker level are available throughout the period. Second, Chilean labour law draws a sharp legal distinction between open ended and fixed term contracts, and contract composition is directly observable in the data, a margin on which the existing climate and labour literature has had less visibility. Third, agriculture remains strategically important for exports and relies heavily on seasonal wage labour, in which women are disproportionately represented in harvesting and agro-industrial tasks (Barrientos and Visser, 2012; Food and Agriculture Organization, 2023). In addition, many local services are tightly linked to agriculture, so drought can widen gender gaps both directly and through spillovers to manufacturing and services (Ngai and Petrongolo, 2017; Pérez-Silva et al., 2023).

I estimate a stacked event study around 24 severe drought onsets following Baker et al. (2022), with cohort specific fixed effects and clean control groups. The design exploits provincial cohorts for direct agricultural effects and regional cohorts for cross-sector spillovers, with the choice driven by the unit at which each effect operates.

Three results stand out. First, within agriculture, hours among workers who remain employed change little at drought onset, but women's agricultural employment share falls by 0.8 percentage points on impact and remains depressed by 1.4 percentage points one quarter later. Men's employment share shows no significant adjustment in this window. The asymmetry is consistent with women bearing the participation margin burden through

non-renewal of seasonal matches, while continuing workers experience little change in hours. Second, severe drought onsets generate sizeable and persistent reductions in non-agricultural hours in regions with higher baseline agricultural dependence. Services hours for men decline by 3–5 hours per week and remain below baseline through ten quarters after onset; the response for women is similarly persistent. Manufacturing hours also fall, with effects concentrated at medium horizons. Third, drought exposure at the origin is associated with higher interregional migration, but this relationship is weaker in more diversified origins. Higher out migration does not buffer women's earnings, the interaction between out migration and a female indicator is negative and significant.

This paper makes three contributions. First, it provides evidence that in a diversified labour market, drought widens gender gaps primarily through participation and job quality margins, even when hours among employed workers adjust little. This extends the drought and gender literature (Afridi et al., 2022; Mahajan, 2017) by documenting contract based adjustment channels that are central in seasonal agricultural labour markets but rarely observed in existing work. Second, it documents that indirect spillovers from agriculture to manufacturing and services can be large and persistent, linking climate shocks to non-agricultural labour outcomes through local demand and production linkage channels. This complements evidence on agricultural and non-agricultural linkages (Emerick, 2018; Blakeslee et al., 2020) by tracing dynamic, gender differentiated spillovers in a diversified setting. Third, it provides evidence that local economic structure shapes spatial adjustment, with diversification attenuating migration induced by droughts and mitigating negative effects on wages for women who remain, extending work on climate and migration (Feng et al., 2010; Jessoe et al., 2018; Nawrotzki et al., 2017).

The remainder of the paper proceeds as follows. Section 2 describes Chile's agricultural labour markets and climate exposure. Section 3 develops the theoretical framework. Section 4 presents the data. Section 5 details the empirical strategy. Section 6 reports the main results. Section 7 discusses robustness. Section 8 concludes.

2 Background

2.1 Droughts in Chile

Chile has experienced an unprecedented sequence of rainfall deficits since 2010, commonly referred to as the megadrought (Garreaud et al., 2017). Central and south central Chile have received 25–45% less precipitation than historical averages, with cumulative deficits unprecedented in reconstructed climate records (Garreaud et al., 2020), and the resulting

reductions in river flows and aquifer levels have imposed binding constraints on irrigated and rainfed agriculture across multiple regions (Dirección General de Aguas, 2025). The severity of the shock varies substantially across the country's sixteen administrative regions. Northern regions are structurally arid with limited agriculture, and southern regions are wetter and have experienced smaller relative deficits, while the sharpest drought episodes have been concentrated in the agricultural heartland (regions V–VIII). Within this heartland there is also substantial quarter to quarter variation in the timing and intensity of moisture deficits (Figures 1 and 2), and it is this combination of spatial and temporal heterogeneity that the empirical strategy exploits.

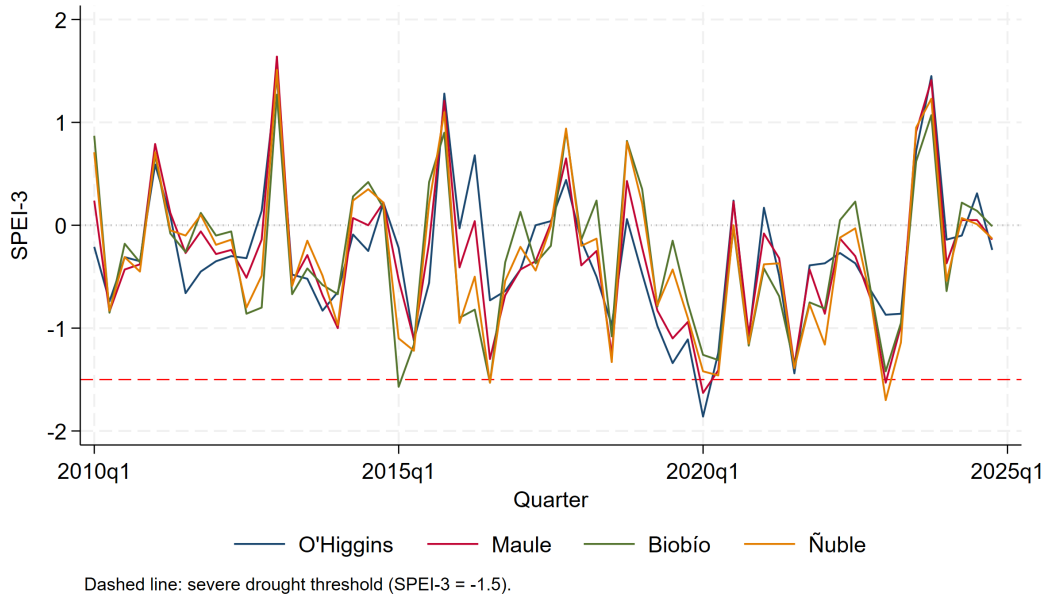


Figure 1: SPEI-3 time series for selected regions

Notes: Quarterly SPEI-3 values for selected Chilean regions, 2010–2024. The dashed horizontal line marks the severe drought threshold (−1.5).

Chile's Mediterranean climate reinforces this variation through a pronounced dry season (November–March) and wet season (May–September), to which agricultural production cycles are tightly linked. Planting is concentrated in austral spring, irrigation intensive growth occurs in summer, and harvest takes place in late summer and autumn, so drought episodes that coincide with sensitive points in the production cycle have disproportionate effects on yields and on the seasonal labour matches that scale with realised volumes. A further feature of Chile's exposure is the heavy dependence of agriculture on surface water for irrigation. Rainfed agriculture in southern regions is directly exposed to precipitation deficits, while irrigated areas in central Chile are exposed through reduced river flows and aquifer depletion, though with a lag (Dirección General de Aguas, 2025; Oficina de Estudios y Políticas

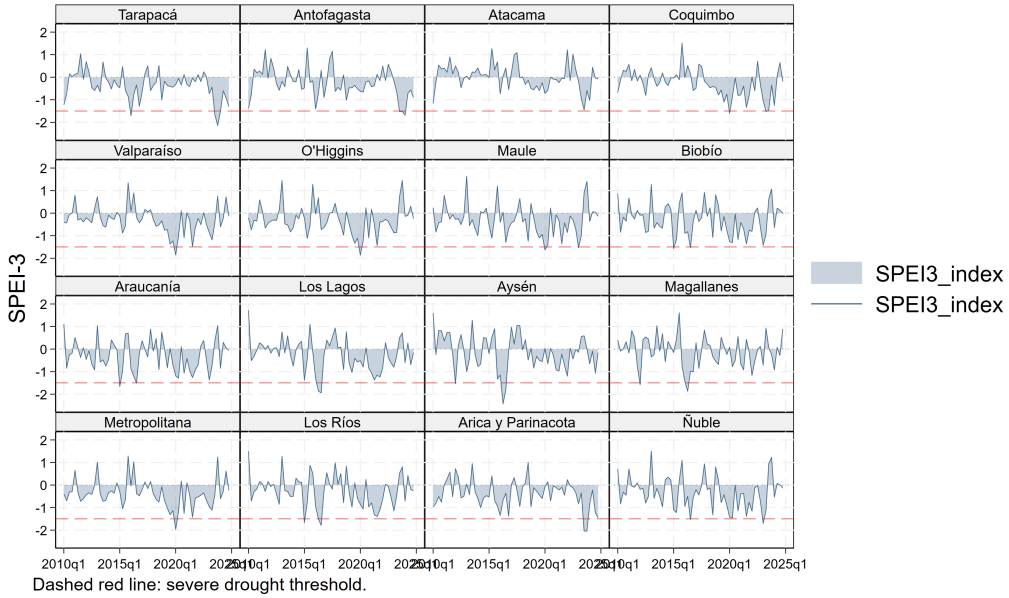


Figure 2: SPEI-3 across all regions

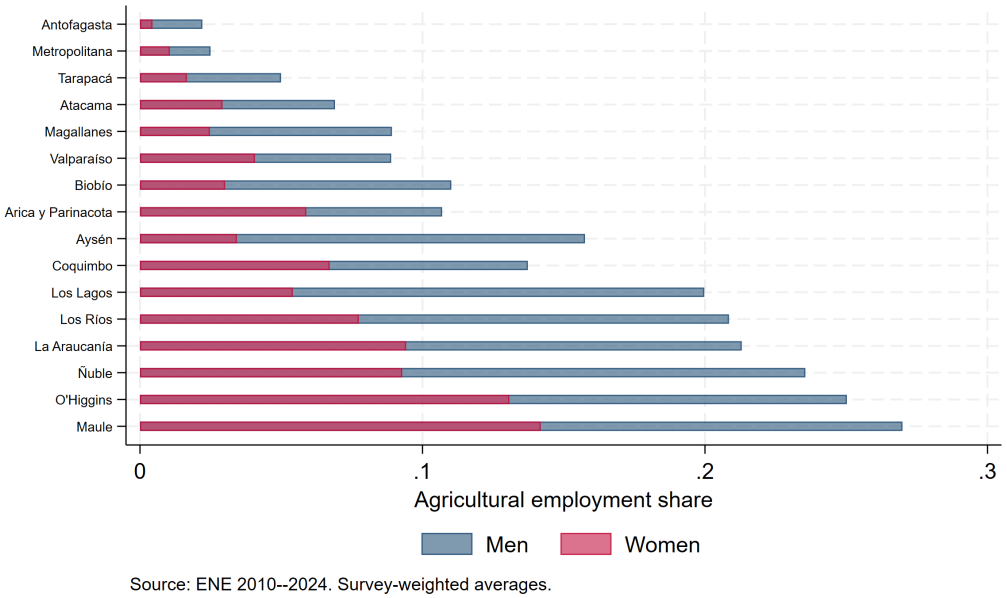
Notes: Quarterly SPEI-3 values for all sixteen Chilean administrative regions, 2010–2024.

Agrarias, 2025). The same national drought episode therefore transmits to agricultural production through distinct hydrological channels across regions, providing additional variation in the effective intensity and timing of the shock.

2.2 Labour market composition in Chile

At the national level, agriculture accounts for approximately 9% of total employment, manufacturing for 10%, and services for roughly 65%, with the remaining 16% in other sectors (ENE averages, 2010–2024). These national figures mask pronounced regional variation. In the central agricultural regions (O'Higgins, Maule, Ñuble), agriculture's employment share reaches 15–25%, while in metropolitan Santiago it falls below 3% (Figure 3). This regional variation in agricultural dependence is central to understanding the potential spillover effects, because it determines the strength of local demand and production linkages through which drought shocks propagate from agriculture into manufacturing and services.

Gender gaps in Chile's labour market are well documented. Female labour force participation reached 60.5%, compared with 77.6% for men in 2023 (ages 15–64) (Organisation for Economic Co-operation and Development, 2025). Within agriculture, women are concentrated in seasonal harvesting and agro-industrial tasks, such as fruit packing and quality sorting, organised through limited term contracts. These workers, known as temporeras, are a defining feature of Chilean agricultural labour markets (Food and Agriculture Organiza-



Source: ENE 2010--2024. Survey-weighted averages.

Figure 3: Agricultural employment share by region

Notes: Average agricultural employment share by region, 2010–2012. Regions above the median are classified as high agricultural dependence in the spillover analysis.

tion, 2023). Outside agriculture, women are disproportionately represented in commerce, hospitality, education, and health, sectors with varying exposure to local demand conditions and institutional stability (Figure 4). This sectoral segmentation is the lens through which the gendered consequences of drought emerge in a diversified economy. Women's concentration in seasonal agricultural tasks exposes them directly to non-renewal of fixed term matches when yields fall, while their concentration in consumer facing services exposes them indirectly to demand side contractions when farm income declines.

Chilean labour law reinforces this segmentation. Open ended contracts carry significant dismissal costs, including severance of one month's pay per year of service, while fixed term contracts expire automatically, making non-renewal a low cost adjustment margin. The prevalence of fixed term contracting in agriculture therefore means that drought adjustment operates primarily through contract non-renewal rather than through proportional reductions in hours or wages among continuing workers, and women's disproportionate exposure to this margin is what links the institutional structure to the gendered patterns documented in this paper. This, combined with regional variation in agricultural dependence and the local input and output linkages that connect agriculture to manufacturing and services, these features create the conditions for the cross sector spillover effects that the empirical analysis identifies.

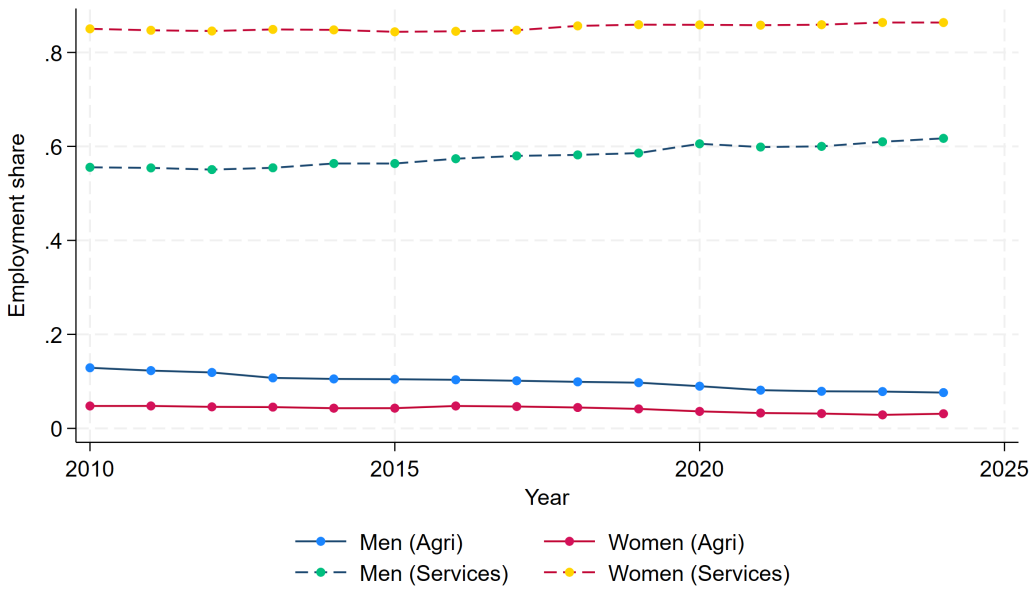


Figure 4: Sectoral employment shares by gender

Notes: Employment shares in agriculture, manufacturing, and services by gender, ENE averages 2010–2024.

3 Theoretical Framework

A key consideration is through which channels a severe drought reshapes labour market outcomes in a diversified economy, and how the burden of adjustment falls differently on women and men. The answer depends on three interconnected channels of transmission, direct displacement within agriculture, cross sector spillovers through local demand and production linkages, and spatial reallocation through interregional migration, whose relative strength varies with the structure of the local economy. Gender differences operate as a cross cutting modifier whose intensity depends on women's and men's position within agricultural and non-agricultural labour markets.

Consider a regional economy that is small relative to national markets. Production takes place in three broad sectors, namely agriculture, which is directly exposed to climatic risk and produces tradable output; manufacturing, which includes tradable activities and, in many regions, activities connected to agricultural value chains such as processing; and services, which include locally oriented non-tradables such as commerce and hospitality alongside branches tied more closely to production volumes and firm activity, such as transport, communications, finance, and business services.[1] Regions differ along two dimensions that shape the transmission of drought shocks. First, baseline agricultural dependence, the extent to

[1]Education, health, and other public services may be less sensitive in the short run because staffing and budgets are shaped by institutional rules and public provision.

which local employment and income rely on agriculture and linked activities. Second, local economic diversification, the breadth of employment opportunities outside agriculture within the same region. A more diversified local economy offers more scope for within region reallocation, dampening both direct and indirect effects. These two dimensions interact with the contract structure of employment. Some dependent employment is organised through open ended contracts that carry significant adjustment costs, while other employment operates through limited term contracts and seasonal arrangements that expire automatically, making non-renewal a low cost margin of adjustment.

The first channel is direct displacement within agriculture. A severe drought reduces agricultural productivity by limiting water availability and increasing crop stress, with immediate consequences for water dependent activities and changes in production timing. When drought coincides with sensitive stages of the production cycle, yields fall sharply; when drought persists, it alters planting decisions, crop composition, and expectations about future conditions. Lower productivity reduces the return to labour in activities that are complementary to realised output (Jayachandran, 2006). Labour intensive stages that scale with volumes can contract quickly when volumes decline or when harvest windows shorten, while other activities may respond with delay as producers revise input use and production plans under prolonged stress. Total labour input in agriculture can adjust through three margins, namely hours per worker, the number of workers employed, and the composition of employment relationships. In seasonal agricultural labour markets, the latter two margins are particularly important. Many tasks are organised around discrete stages and workdays, so hours recorded for those who remain employed can change little even when total labour input falls, and adjustment operates through the disappearance of marginal matches rather than through proportional hour cuts for all workers. Limited term and seasonal arrangements allow firms to scale labour input by not renewing matches, reducing seasonal hiring, or delaying recruitment during peak periods, while open ended matches are more costly to adjust. Job stability is therefore itself economically relevant, since reliance on limited term contracting can increase when uncertainty rises while also raising income volatility and insecurity for workers, so drought can worsen labour outcomes through job quality even if wages or hours among continuing workers do not fall sharply on impact.

The second channel is cross sector spillovers. Drought impacts propagate beyond agriculture through two distinct pathways. The first is a local income channel. A fall in farm revenues and agricultural employment reduces local disposable income and increases uncertainty, depressing consumption of locally oriented services. This pathway should be strongest in consumer facing non-tradables, especially commerce and hospitality, where activity depends on local spending and cannot be readily redirected to outside markets in the short

run. Firms in these sectors adjust by reducing labour input, which shows up as lower hours among the employed and weaker employment. The second is a production linkage channel. Lower agricultural volumes reduce activity in upstream and downstream sectors tied to agriculture, including processing and logistics, transmitting drought effects to manufacturing and to services connected to volumes and firm activity such as transport, communications, finance, and business services (Bastos et al., 2013; Branco and Féres, 2021). Timing can differ from the demand channel because volume reductions take time to pass through processing, storage, and transport networks. Both pathways scale with baseline agricultural dependence, since when agriculture is a larger share of local income and activity, the income shock and the linkage shock are mechanically larger, so spillovers should be more pronounced in agriculture dependent regions. Where local economies are more diversified, other sources of income and employment can partially offset agricultural losses, dampening spillovers. Within services, heterogeneity across sub-sectors is informative about which pathway dominates. Larger effects in commerce and hospitality are consistent with the income channel, responses in transport, communications, and business services are consistent with the linkage channel, and muted responses in public services, education, and health are consistent with institutional rigidity and public provision.

The third channel is interregional migration. When direct displacement within agriculture and indirect contraction in linked sectors exhaust local reallocation possibilities, spatial adjustment becomes an additional margin. Drought can increase out migration from affected origin regions when local earnings opportunities deteriorate and recovery is perceived to be slow or uncertain (Feng et al., 2010; Hornbeck, 2012; Nawrotzki et al., 2017). Migration is costly and often reflects household level decisions, so it is typically a residual response, activated only when within region reallocation has reached its limits. Local diversification reduces the need for spatial adjustment. A more diversified origin region offers more outside options outside agriculture, facilitating within region transitions and attenuating drought induced out migration. Out migration does not mechanically improve outcomes for those who remain. If drought depresses local labour demand beyond agriculture through both the demand and linkage channels, out migration can coincide with weak earnings and weak hours for stayers. This is especially relevant when the contraction is concentrated in locally oriented services where reallocation options are limited, and it implies that the welfare consequences of migration depend not only on who moves but on the breadth of the local contraction faced by those who stay.

A key implication of these three channels is that gender differences arise endogenously from where women and men sit within the labour market structure rather than being a separate channel of their own. Two considerations are particularly relevant. First, exposure

to flexible margins. Seasonal agriculture and agro-industry rely heavily on limited term contracting, and if one gender is more represented in this form of employment, then first round adjustment, that is, non-renewal and reduced seasonal hiring, translates into larger immediate employment losses for that group. This mechanism does not require a specific assignment of women and men to particular tasks; it operates through differential exposure to contract types (Mahajan, 2017; Eastin, 2018; Afridi et al., 2022). Second, outside options and reallocation. The local composition of vacancies matters because outside options can differ by gender if women and men are concentrated in different parts of the non-agricultural economy or in job forms with different stability. When outside options differ, the same shock generates different trajectories, with some workers reallocating across sectors and others experiencing longer non-employment spells or transitions into less stable jobs. These two considerations imply that gender gaps can widen through participation and job stability margins even when hours among those who remain employed adjust little, and that women whose concentration in consumer facing services exposes them to demand side spillovers may not benefit from the spatial adjustment that out migration provides for men in more tradable sectors.

A stylised one period model formalising these channels and yielding the comparative statics that map directly to the empirical specifications below is presented in Appendix 8.

The framework yields the following predictions:

(i) Within agriculture, adjustment should occur primarily through participation and job stability margins, not through hours among continuing workers. Gender differences may arise in timing, with earlier responses in women's employment shares and later responses in men's contract composition.

(ii) Spillovers to non-agricultural hours should be larger in regions with higher baseline agricultural dependence, where income and production linkage channels are mechanically stronger.

(iii) Within services, heterogeneity should be informative about propagation channels. Larger effects in commerce and hospitality are consistent with the income channel, while responses in transport, communications, and business services are consistent with the linkage channel.

(iv) Migration should respond to drought but be attenuated by local diversification, and out migration should not mechanically improve outcomes for women who stay.

4 Data and Descriptive Statistics

4.1 Data

The main labour dataset is the National Employment Survey (ENE), produced by Chile's National Statistics Institute (INE). ENE is a continuous labour force survey with national coverage and a probabilistic, stratified sampling design that is representative at the regional level. The survey covers the working age population (15 years and older) and collects detailed information on employment status, occupation, industry, hours worked, and contract type. ENE has a rotating panel structure in which dwellings are interviewed for up to six rounds over eighteen months, with interviews spaced at roughly three month intervals. I align the data at the quarterly frequency with drought exposure, using ENE expansion factors (survey weights) in estimation. The estimation sample includes all individual quarter observations from the ENE microdata for 2010Q1 through 2024Q4, restricted to individuals aged 15–64 and assigned to one of Chile's sixteen administrative regions based on their region of residence at the time of interview. The final sample contains approximately 5.7 million individual quarter observations, with roughly equal representation of men and women.

For the analysis I use four sets of outcome variables. The primary intensive margin outcome is effective hours worked, defined as weekly hours worked in the reference week. The extensive margin is captured by sectoral employment shares, constructed from the industry classification of the main job. I focus on three broad sectors, namely agriculture, manufacturing, and services, and also study a four way decomposition of services into consumer market services (commerce and hospitality), transport and communications, producer and business services (finance, real estate, professional services), and public and social services (education, health, public administration). This partition differentiates branches plausibly more exposed to local demand from those partially stabilised by public provision, and maps directly to the distinction between the income channel and the production linkage channel introduced in Section 3. To study job quality, I use contract type information for dependent employees, measuring the share of dependent agricultural workers holding limited term contracts and related indicators of contract duration and seasonality. Earnings outcomes are measured using the Complementary Income Survey (ESI), an income module fielded within ENE once per year in the October–December quarter. I construct an hourly wage proxy as monthly labour income divided by effective hours worked. Given that ESI is annual, earnings results are treated as complementary evidence rather than as a primary outcome.

Internal migration is measured using the ENE migration module, which records the region of residence five years earlier. An individual is classified as an interregional migrant if their current region differs from the region reported five years ago. I construct origin region

migration rates and examine how drought exposure at the origin relates to out migration. To capture local economic structure, I compute a diversification index based on the sectoral distribution of employment: $DIV_r = 1 - HHI_r$, where HHI_r is the Herfindahl-Hirschman concentration index of sectoral employment shares in region r. Higher values indicate a more balanced employment structure across sectors. This index serves as the key moderator in the migration analysis, capturing the availability of within region alternatives that attenuates the need for spatial adjustment.

Drought exposure is measured using the Standardised Precipitation Evapotranspiration Index (SPEI) derived from the CR2MET gridded climate dataset produced by the Center for Climate and Resilience Research (CR2, Universidad de Chile). SPEI summarises moisture conditions by combining precipitation with potential evapotranspiration, which captures the role of temperature in water stress. I use a three month accumulation window (SPEI-3), which aligns naturally with quarterly labour outcomes and captures the cumulative moisture deficit relevant for agricultural water availability. SPEI is standardised so that values are comparable across locations and seasons; negative values indicate drier than normal conditions. Following standard thresholds in the climatological literature (McKee et al., 1993; Vicente-Serrano et al., 2010), I define severe drought as SPEI-3 ≤ -1.5 and extreme drought as SPEI-3 ≤ -2.0. The main treatment is a drought onset, defined as the first quarter in which SPEI-3 crosses the severe threshold after being above it in the previous quarter. This threshold crossing definition isolates moments when climatic stress becomes sufficiently intense to affect production decisions and labour allocation, and its staggered nature across regions generates the identifying variation that I exploit (Figure 5).

The variable for the spillover analysis is baseline agricultural dependence, measured as the agricultural employment share at the region level averaged over 2010–2012. This pre-period measure captures how central agriculture is to local demand and production linkages, and its interaction with event time indicators identifies whether drought induced contractions in non-agricultural sectors are larger in regions where agriculture accounts for a greater share of local economic activity. The variable for the migration analysis is the diversification index described above, which captures the breadth of employment opportunities outside agriculture within the same region.

4.2 Descriptive statistics

Table 1 presents summary statistics for the estimation sample. The sample is roughly balanced by gender, with women constituting 50.2% of observations, and the average age is 38 years with similar distributions across genders. About 62% of the sample is employed, with

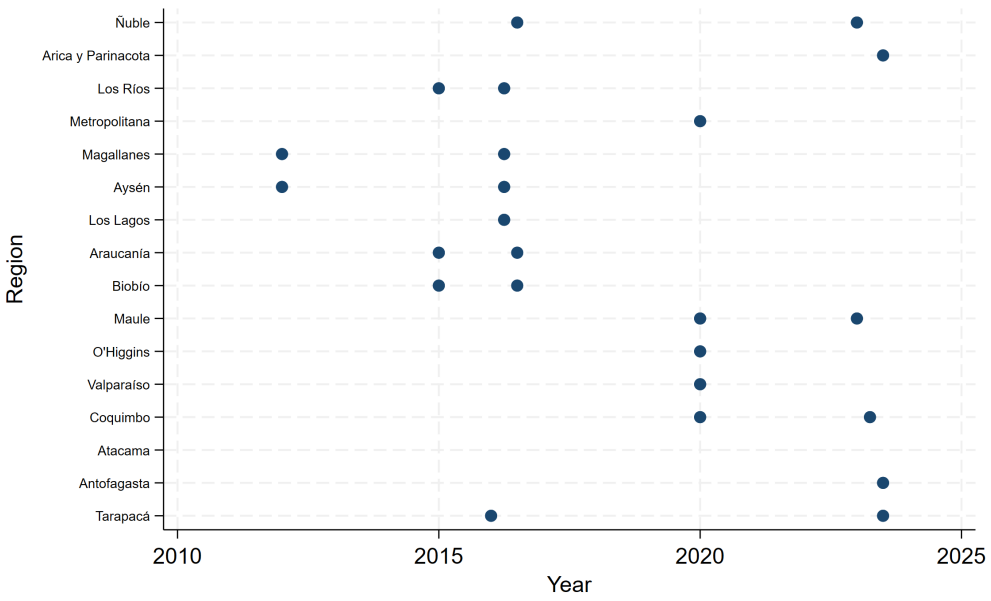


Figure 5: Timing of drought onsets across regions

Notes: Each marker indicates a drought onset (first quarter with SPEI-3 ≤ -1.5 following a quarter above the threshold) by region and calendar quarter.

a substantial gender gap (72.3% for men and 51.8% for women) that is itself informative about the baseline conditions through which drought shocks are filtered. Among the employed, men work an average of 42.3 effective weekly hours compared to 35.4 for women. The sectoral distribution reveals the gender segmentation central to this paper, with 9.9% of employed men working in agriculture compared to 4.0% of women, while 85.3% of employed women are in services compared to 58.0% of men. This concentration of women in services, and particularly in consumer facing branches, is what makes demand side spillovers from agricultural income contractions a first order channel for gender differentiated adjustment. At the climate margin, the mean SPEI-3 index is -0.26 at the region and quarter level, reflecting the drier than normal conditions that have prevailed during the megadrought, and the drought indicator (SPEI-3 ≤ -1.5) is activated in 3.3% of region quarters, providing sufficient variation in onset timing across regions to identify the stacked event study design.

The gender gap in effective hours documented in Table 1 is a stable feature of the Chilean labour market over the sample period rather than a recent development. Figure 6 plots the time series of average effective weekly hours among employed workers separately for men and women. The level difference of roughly 7 hours per week persists throughout 2010–2024, with no narrowing over time. This gap reflects the combination of women's concentration in part time, seasonal, and consumer facing service positions and men's concentration in full time agricultural, manufacturing, and supervisory roles, and it is the baseline against which

Table 1: Summary Statistics

	Men		Women	
	Mean	SD	Mean	SD
Panel A: Individual characteristics				
Age	37.799	14.059	38.388	14.186
Primary education	0.165	0.372	0.161	0.367
Secondary education	0.470	0.499	0.469	0.499
Tertiary education	0.350	0.477	0.359	0.480
Employed	0.723	0.448	0.518	0.500
Panel B: Labour market outcomes (employed)				
Effective weekly hours	42.346	40.183	35.352	35.413
Employed in agriculture	0.099	0.299	0.040	0.197
Employed in manufacturing	0.321	0.467	0.106	0.308
Employed in services	0.580	0.494	0.853	0.354
Panel C: Climate variables (region-quarter level)				
	Mean		SD	
SPEI-3 index	−0.261		0.666	
Drought indicator	0.033		0.180	

Notes: Panel A reports weighted means and standard deviations for the full individual-quarter sample. Panel B conditions on employed individuals. Panel C is measured at the region-quarter level and does not vary by gender; the gender columns are pooled to display the single mean and standard deviation. Data: ENE 2010–2024, SPEI-3. Weights: ENE survey expansion factors.

the drought induced changes in hours documented in Section 6 should be read.

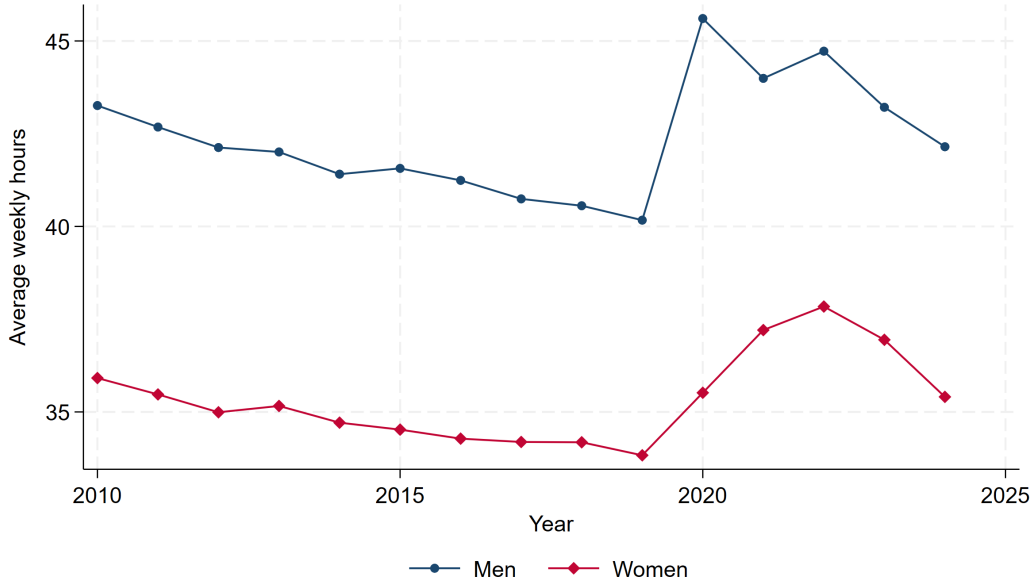


Figure 6: Average effective hours by gender, 2010–2024

Notes: Average weekly effective hours among employed workers by gender. Source: ENE 2010–2024, weighted by survey expansion factors.

5 Empirical Strategy

I estimate the dynamic effects of drought onset using a stacked event study design following Baker et al. (2022). Let i index individuals, r regions, p provinces (with $p \in r$), t calendar quarters, and e drought onset events. A drought onset is defined as the first quarter in which SPEI-3$_r$ crosses the -1.5 threshold after being above it in the previous quarter. The shock is therefore measured at the regional level, since drought is a basin scale meteorological event and finer gridded SPEI within a region carries measurement noise rather than additional climatic signal. I implement a proper stacked design that constructs event specific cohorts with clean control groups. A cohort is defined as a single drought onset event together with all observations of treated and clean control units within the event window of five quarters before and ten quarters after that onset; the panel is then assembled by stacking these cohorts so that each treated unit is compared only to controls observed in the same calendar window, with cohort specific fixed effects absorbing any cohort level confounders. This is necessary because a region can experience multiple drought onsets during the sample period, and because standard two way fixed effects event studies with staggered treatment timing

can produce biased estimates when already treated units serve as controls (Callaway and Sant'Anna, 2021; Sun and Abraham, 2021; Goodman-Bacon, 2021; Borusyak et al., 2024).

For the direct effects analysis I exploit the fact that Chile's 16 regions partition into 54 provinces. Although drought onset is regional, I construct cohorts, fixed effects, and clusters at the provincial level. Each province in a treated region is a treated unit, and clean controls are provinces in regions with no onset within $[t_e - 5, t_e + 10]$. This raises the number of clusters from 16 to 54, mitigating the small cluster concerns of Cameron et al. (2008) and Abadie et al. (2023), and supports finer demographic fixed effects (cohort × province × age × education × season). The stacking procedure identifies 68 drought onset cohorts under this provincial scheme.

For the spillover specification below I retain the regional structure: 24 cohorts, 16 clusters, and small cluster inference complemented by wild cluster bootstrap. The reason is that the heterogeneity dimension that identifies spillovers, baseline agricultural dependence, is itself a regional construct, since production and demand linkages connecting agriculture to non-agricultural sectors operate across provinces within a region rather than within a single province. The estimating equation for the direct effects specification is:

$$y_{ipte} = \sum_{\substack{k=-5 \\ k\neq -1}}^{10} \beta_k \cdot \mathbf{1}[t - t_e = k] + \alpha_{e,t} + \gamma_{e,p\times a\times d\times q} + \varepsilon_{ipte} \tag{1}$$

where y_{ipte} is the outcome for individual i in province p, calendar quarter t, within the window of event e. The β_k coefficients trace the dynamic response at each event time k relative to the omitted period $k = -1$. The term $\alpha_{e,t}$ denotes cohort × calendar quarter fixed effects that absorb aggregate time shocks within each cohort. The term $\gamma_{e,p\times a\times d\times q}$ denotes cohort × province × age group × education group × quarter of year fixed effects, which flexibly control for province specific seasonal patterns and demographic composition within each cohort. All specifications use ENE survey expansion factors as analytic weights and are estimated separately by gender (Correia, 2016). Standard errors are clustered at the province level (54 clusters). The coefficients of interest are $\{\beta_k\}_{k=0}^{10}$, which capture the post-onset dynamics. The lead coefficients $\{\beta_k\}_{k=-5}^{-2}$ serve as a pre trends diagnostic. Under the identifying assumption that drought onset timing is conditionally exogenous given the fixed effects, these lead coefficients should be close to zero.

To estimate cross sector spillovers, I interact the event time indicators with an indicator for whether the region has above median baseline agricultural dependence, measured as the agricultural employment share averaged over 2010–2012:

$$y_{irte} = \sum_{\substack{k=-5 \\ k\neq -1}}^{10} [\beta_k \cdot \mathbf{1}[k] + \delta_k \cdot \mathbf{1}[k] \times \text{HighAg}_r] + \alpha_{e,t} + \gamma_{e,r\times a\times d\times q} + \varepsilon_{irte} \tag{2}$$

where HighAg_r is an indicator equal to one for regions with above median agricultural employment shares in 2010–2012. The interaction coefficients $\{\delta_k\}$ capture how outcomes in manufacturing and services evolve differentially in more agriculture dependent regions around drought onset. Given that the stacked design includes cohort specific fixed effects and compares high and low dependence regions within the same cohort and quarter, it controls for national shocks and cohort specific confounders, isolating the differential adjustment in more exposed local economies. Equation (2) is estimated at the regional level: 24 cohorts, cohort × region × age × education × season fixed effects, and standard errors clustered at the region level (16 clusters), complemented by wild cluster bootstrap p-values for the post-onset interaction coefficients.

For migration, I estimate a linear probability model at the individual level that exploits cross sectional variation in drought exposure at the origin region together with the diversification of the local economy:

$$\text{Migrant}_{irt} = \beta_1 \cdot \text{Drought}_{r_0 t} + \beta_2 \cdot \text{DIV}_{r_0} + \beta_3 \cdot (\text{Drought}_{r_0 t} \times \text{DIV}_{r_0}) + X_{it}'\pi + \mu_{r_0} + \lambda_t + \varepsilon_{irt} \tag{3}$$

where r_0 denotes the origin region, $\text{Drought}_{r_0 t}$ is an indicator for SPEI-3 ≤ -1 in the origin, DIV_{r_0} is the diversification index, and X_{it} includes age (quadratic) and education dummies. The migration specification uses a less restrictive threshold (SPEI-3 ≤ -1) than the direct effects (≤ -1.5) because severe drought events are too rare at the individual annual level to support the interaction model. Origin region fixed effects μ_{r_0} absorb time invariant differences across origins, and year fixed effects λ_t absorb common shocks. The coefficient β_3 tests whether diversification attenuates drought induced out migration.

I complement this specification with a stayer earnings regression that tests whether out migration translates into higher local earnings for women who remain. Restricting the sample to employed individuals whose current region of residence coincides with their origin region (stayers), I estimate:

$$\log w_{irt} = \alpha_1 \cdot \text{OutMig}_{r_0 t} + \alpha_2 \cdot \text{Female}_i + \alpha_3 \cdot (\text{OutMig}_{r_0 t} \times \text{Female}_i) + Z_{it}'\zeta + \mu_{r_0} + \lambda_t + \nu_{irt} \tag{4}$$

where $\log w_{irt}$ is the log hourly wage proxy from the ESI module, $\text{OutMig}_{r_0 t}$ is the out

migration rate of the origin region in year t, Female$_i$ is a female indicator, and Z_{it} contains the same age and education controls as in equation (3). Origin region and year fixed effects again absorb time invariant origin heterogeneity and aggregate trends. The coefficient of interest is α_3, which captures the differential earnings response of women relative to men as origin out migration rises. A standard tightening prediction would imply $\alpha_1 > 0$ for stayers in general, while a coefficient $\alpha_3 < 0$ would indicate that the earnings benefits of out migration accrue disproportionately to men, leaving women in more demand exposed segments behind.

The identifying assumption is that, conditional on the cohort specific fixed effects in equations (1) and (2), the timing of severe drought onset is orthogonal to unobserved determinants of labour market outcomes. Four features support this assumption. First, SPEI-3 captures exogenous climatic variation driven by precipitation and temperature anomalies, which are not determined by local economic decisions. Second, the stacked design with clean control groups avoids contamination from already treated units acting as controls, a well documented concern in standard two way fixed effects event studies with staggered treatment (Callaway and Sant'Anna, 2021; Sun and Abraham, 2021; Goodman-Bacon, 2021; Borusyak et al., 2024), since each treated unit is compared only to genuinely untreated regions within the same time frame. Third, the lead coefficients $\{\beta_k\}_{k=-5}^{-2}$ provide a transparent pre trends diagnostic, and I report joint F-tests for the null that all leads are zero alongside the individual lead coefficients in the event study figures. Fourth, I provide a pre-treatment covariate balance table (Table 15) comparing demographic and employment characteristics between treated and clean control regions in the pre-event period, and complement it with a placebo test using public sector employment as an outcome that should not respond to drought through the agricultural channel (Table 16).

The number of clusters depends on the specification. The direct effects specifications use 54 provincial clusters, well within the range where the conventional asymptotic approximation of cluster robust standard errors performs adequately (Cameron and Miller, 2015). The spillover specifications retain 16 regional clusters; with few clusters, conventional cluster robust standard errors may under-reject the null hypothesis (Cameron et al., 2008; Abadie et al., 2023), so I complement clustered inference with wild cluster bootstrap (WCB) p-values for key post-treatment coefficients (Cameron et al., 2008; Roodman et al., 2019). In addition, throughout the analysis I report the full event time coefficient paths, summarise dynamics into post-onset blocks (quarters 0–2, 3–6, 7–10) with joint Wald tests, and implement leave-one-out robustness checks that sequentially drop each cluster to verify the results are not driven by any single one.

6 Results

6.1 Direct effects within agriculture

If drought operates as a negative productivity shock in a seasonal labour market, the standard prediction is that adjustment should occur primarily through participation margins, that is, the non-renewal of seasonal matches and exit from employment, rather than through hours or wage cuts among continuing workers. The intensive margin estimates confirm this prediction. Average effective hours among agricultural workers who remain employed change little at drought onset for either gender (Table 2; Figure 7). The point estimates are small and statistically insignificant at all post-onset horizons, consistent with the structure of Chilean agricultural employment, where open ended contracts carry severance costs that discourage hour reductions while seasonal contracts either continue at their agreed terms or are simply not renewed.

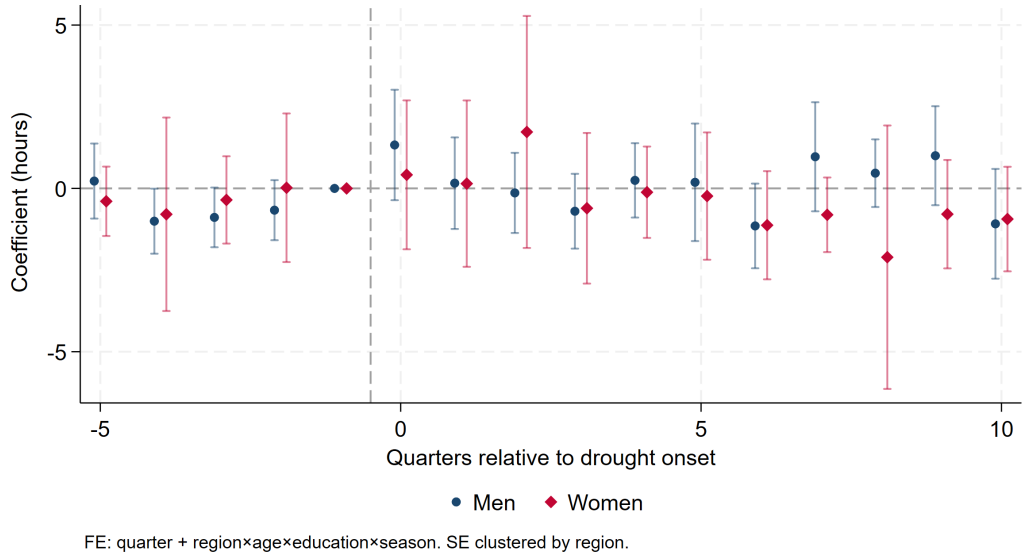


Figure 7: Event study estimates: Effective hours in agriculture

Notes: Coefficients β_k from equation (1) estimated separately by gender. Vertical bars show 95% confidence intervals based on cluster robust standard errors (54 provincial clusters). Omitted period: $k = -1$.

The extensive margin tells a different story, and one that diverges sharply by gender (Table 3). Women's agricultural employment share drops by 0.8 percentage points at onset ($p < 0.05$), a decline equivalent to roughly 20% of women's baseline agricultural employment share of 4.0%, and falls further at $k = +1$ (-1.4 pp, $p < 0.01$), with significant negative effects persisting through $k = +5$. The immediacy of this response is consistent with the non-renewal of seasonal contracts that coincide with the planting to harvest cycle. When drought reduces expected yields, employers choose not to hire or re-engage temporeras for

Table 2: Direct Effects in Agriculture — Intensive Margin

	Men	Women
$k = -5$	-0.460	-1.887
	(0.847)	(2.450)
$k = -4$	-1.085	2.221
	(1.003)	(1.989)
$k = -3$	-1.701	2.428
	(1.103)	(2.058)
$k = -2$	1.157	-0.0388
	(0.776)	(1.471)
$k = 0$	1.295	4.603
	(1.358)	(3.531)
$k = +1$	-0.799	3.572
	(1.210)	(2.515)
$k = +2$	2.178*	4.235
	(1.274)	(4.424)
$k = +3$	-1.752*	-2.364
	(1.010)	(1.854)
$k = +4$	-0.364	3.561
	(0.839)	(3.089)
$k = +5$	-0.547	2.497
	(1.023)	(2.010)
$k = +6$	0.728	-1.745
	(1.106)	(1.940)
$k = +7$	0.128	-2.516
	(1.134)	(2.226)
Observations	2,147,866	707,623
R^2	0.079	0.157

Notes: Coefficients β_k from equation (1) estimated separately by gender. Dependent variable: weekly effective hours among workers employed in agriculture. Standard errors clustered at the province level (54 clusters) in parentheses; drought onset is defined at the regional level (see Section 5). Omitted period: $k = -1$. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

tasks such as fruit picking, packing, and quality sorting. Men's agricultural employment share, by contrast, shows no statistically significant adjustment in the post-onset window, consistent with their greater representation in open ended and supervisory positions that are not subject to immediate non-renewal.

Table 3: Direct Effects in Agriculture — Extensive Margin (Agricultural Employment Share)

	Men	Women
$k = -5$	-0.00437	-0.00508*
	(0.00426)	(0.00288)
$k = -4$	-0.00810	-0.00666
	(0.00659)	(0.00622)
$k = -3$	-0.00413	-0.00868
	(0.00546)	(0.00562)
$k = -2$	0.00127	-0.00258
	(0.00654)	(0.00431)
$k = 0$	-0.00938	-0.00821**
	(0.00593)	(0.00406)
$k = +1$	-0.0112	-0.0141***
	(0.00692)	(0.00387)
$k = +2$	0.000506	-0.000374
	(0.00535)	(0.00352)
$k = +3$	-0.00331	-0.00776**
	(0.00420)	(0.00327)
$k = +4$	-0.00652	-0.000682
	(0.00523)	(0.00414)
$k = +5$	-0.00706	-0.00624**
	(0.00442)	(0.00290)
$k = +6$	-0.00183	0.000267
	(0.00458)	(0.00316)
$k = +7$	-0.00129	-0.00439
	(0.00662)	(0.00423)
Observations	15,114,845	11,959,530
R^2	0.196	0.122

Notes: Coefficients β_k from equation (1) estimated separately by gender. Dependent variable: indicator for employment in agriculture, expressed as a share. Standard errors clustered at the province level (54 clusters) in parentheses; drought onset is defined at the regional level (see Section 5). Omitted period: $k = -1$. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Contract composition data are consistent with the participation margin interpretation (Table 4). Among dependent male agricultural workers, the share holding limited term con-

tracts shows no statistically significant change in the post-onset window, with point estimates that are positive but imprecise. Among women, who are already heavily concentrated in limited term and seasonal contracts, the limited term share rises substantially at $k = +2$ and $k = +6$ ($p < 0.01$), reinforcing the picture of an extensive margin adjustment that operates through contract non-renewal rather than through reductions in hours among continuing workers.

Table 4: Mechanism — Limited-Term Contract Share in Agriculture

	Men	Women
$k = -5$	0.0208	0.0429
	(0.0217)	(0.0262)
$k = -4$	-0.0415	0.0621*
	(0.0257)	(0.0335)
$k = -3$	-0.0589**	-0.0259
	(0.0261)	(0.0300)
$k = -2$	-0.0521*	0.123**
	(0.0264)	(0.0461)
$k = 0$	-0.0282	0.000417
	(0.0278)	(0.0212)
$k = +1$	-0.0288	-0.0784
	(0.0259)	(0.0481)
$k = +2$	-0.000207	0.145***
	(0.0290)	(0.0401)
$k = +3$	0.0236	0.0406
	(0.0170)	(0.0370)
$k = +4$	0.0143	0.0415
	(0.0258)	(0.0297)
$k = +5$	-0.0179	-0.0434
	(0.0200)	(0.0385)
$k = +6$	-0.0141	0.132***
	(0.0255)	(0.0474)
$k = +7$	0.0153	0.0192
	(0.0206)	(0.0337)
Observations	1,466,802	472,150
R^2	0.216	0.325

Notes: Coefficients β_k from equation (1) estimated separately by gender. Dependent variable: share of dependent agricultural workers holding limited-term contracts. Sample window is shorter than for other outcomes due to data availability. Standard errors clustered at the province level (54 clusters) in parentheses; drought onset is defined at the regional level (see Section 5). Omitted period: $k = -1$. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Two additional pieces of evidence support this interpretation. First, the ESI earnings module shows no systematic decline in hourly wages in agriculture after drought onset. Together with the stability of hours among survivors, this pattern indicates that the matches that disappear are those at the margin, that is, the seasonal hires that would otherwise have been renewed, so average wages and hours among continuing workers remain largely unchanged even as total agricultural labour input contracts. Second, the gendered timing aligns with findings in Afridi et al. (2022), who document that women in Indian agriculture bear disproportionate adjustment to rainfall shocks. The Chilean results extend this evidence by identifying the specific institutional channel, namely seasonal contract non-renewal, through which this adjustment operates in a diversified labour market with more formalised employment relationships.

6.2 Spillovers to manufacturing and services

A key prediction of the local linkages framework is that drought should reduce labour demand not only in agriculture but also in sectors connected to farm output and farm income. The spillover specification (equation (2)) tests this prediction by comparing labour outcomes in manufacturing and services across regions with above and below median baseline agricultural dependence. If drought reduces local purchasing power and disrupts supply chains, the interaction coefficients δ_k should be negative in post-onset quarters, and the effects should be larger and more persistent in services, which depend heavily on local demand, than in manufacturing, which is more tradable.

The manufacturing response is consistent with production linkage effects that take time to materialise (Table 5; Figure 8). Rather than declining immediately at onset, manufacturing hours in agriculture dependent regions fall by 2–4 hours per week at medium horizons ($k = +4$ to $+8$) for men, with several individual coefficients significant at conventional levels. Relative to a baseline of roughly 42 hours per week, these declines represent a 5–10% reduction in weekly labour input, an economically meaningful contraction, particularly given that it occurs in a sector not directly exposed to drought. For women, the manufacturing response is sharper but less persistent. Hours decline by 4.2–5.3 hours at $k = +3$ to $+4$ ($p < 0.01$), consistent with women's concentration in downstream agro-processing tasks such as packaging, sorting, and light manufacturing, which are directly affected when farm output contracts.

The service sector response is both larger and more persistent, consistent with the local demand channel (Table 6; Figure 9). For men, services hours in agriculture dependent regions decline by 5.5 hours at onset ($p < 0.10$) and remain 2–5 hours below baseline through

Table 5: Spillovers to Manufacturing (Effective Hours × High Agricultural Dependence)

	Men	Women
$k = -5$	-0.0907	-1.192
	(0.640)	(1.349)
$k = -4$	-0.381	-0.988
	(0.481)	(0.599)
$k = -3$	-0.0516	-0.479
	(0.812)	(1.327)
$k = -2$	-1.547*	0.133
	(0.736)	(1.104)
$k = 0$	-3.149	-0.272
	(2.577)	(4.083)
$k = +1$	-2.788	4.253
	(2.112)	(2.761)
$k = +2$	0.838	-0.585
	(1.188)	(2.002)
$k = +3$	-2.107	-4.239***
	(1.874)	(1.074)
$k = +4$	-2.141	-5.298***
	(1.252)	(1.264)
$k = +5$	-2.623**	1.316
	(1.051)	(1.111)
$k = +6$	-3.737**	-0.147
	(1.622)	(0.860)
$k = +7$	-0.858	-6.388**
	(0.678)	(2.874)
$k = +8$	-2.214**	1.169
	(0.835)	(2.100)
$k = +9$	-1.038	-1.108
	(0.601)	(0.994)
$k = +10$	-2.344*	-0.334
	(1.229)	(1.357)
Observations	451,648	112,733
R^2	0.006	0.023

Notes: Interaction coefficients δ_k from equation (2) estimated separately by gender. Dependent variable: weekly effective hours among workers employed in manufacturing. Coefficients capture the differential response of high agricultural dependence regions relative to low dependence regions at each event time. HighAg is defined at the regional level (median split of 2010–2012 agricultural employment shares across the sixteen regions). Standard errors clustered at the region level (16 clusters) in parentheses; wild cluster bootstrap p-values reported in the text for the post-onset coefficients. Omitted period: $k = -1$. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

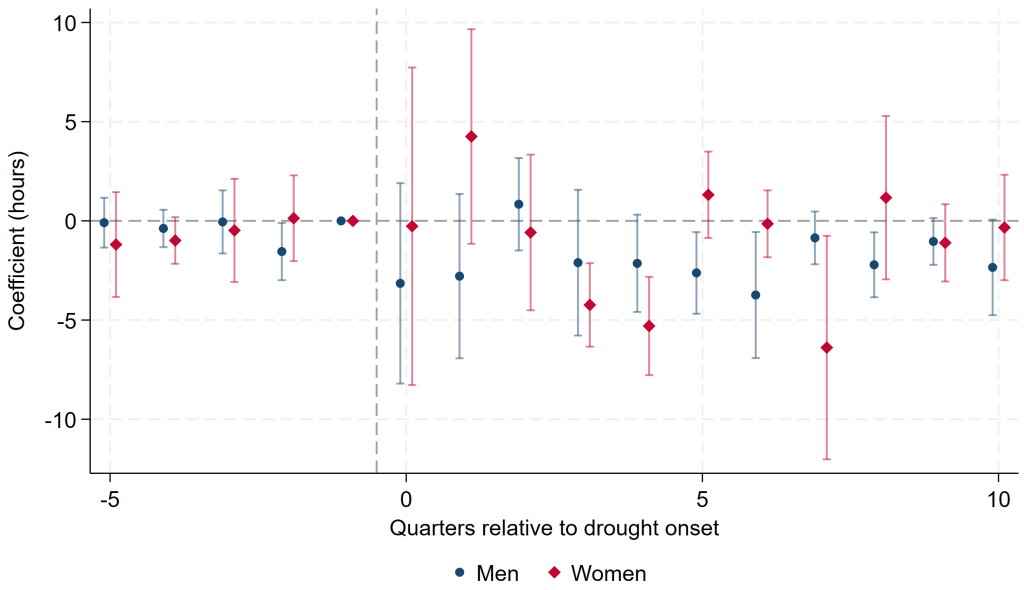


Figure 8: Spillover estimates: Effective hours in manufacturing (× high agricultural dependence)

Notes: Interaction coefficients δ_k from equation (2). Bars show 95% confidence intervals based on cluster robust standard errors (16 clusters). Omitted period: $k = -1$.

$k = +10$, a sustained contraction spanning over two years. For women, the decline materialises at $k = +2$ (-2.1 hours, $p < 0.05$) and persists through $k = +10$ (-2.6 hours, $p < 0.01$). The persistence of these effects distinguishes them from the manufacturing response and is consistent with the mechanism emphasised by Emerick (2018). Agricultural income contractions reduce expenditure on locally produced non-tradable services, generating a demand multiplier that depresses hours and employment in commerce, hospitality, and personal services. In the Chilean context, these are precisely the service branches where women are disproportionately concentrated (Table 1), so the spillover channel amplifies the direct agricultural effect and widens the overall gender gap in drought affected regions.

6.3 Within services heterogeneity

The aggregate services decline documented above could reflect different propagation mechanisms. To discriminate among them, I decompose the services response into four branches with distinct exposure profiles (Tables 7–10; Figures 10 and 11). If spillovers operate primarily through local demand contractions, the largest effects should appear in consumer facing branches such as commerce and hospitality, where revenues depend directly on local purchasing power. If production linkages dominate, transport, communications, and business

Table 6: Spillovers to Services (Effective Hours × High Agricultural Dependence)

	Men	Women
$k = -5$	-0.0542	-0.618
	(0.842)	(0.506)
$k = -4$	0.582	-0.0703
	(0.914)	(0.693)
$k = -3$	-0.865	0.178
	(0.602)	(0.586)
$k = -2$	0.0618	0.0709
	(0.605)	(0.877)
$k = 0$	-5.464	-2.430
	(3.236)	(1.587)
$k = +1$	-0.172	1.553
	(0.993)	(0.895)
$k = +2$	-0.267	-2.094**
	(1.613)	(0.894)
$k = +3$	-2.043	-1.469
	(1.263)	(1.264)
$k = +4$	-3.176**	-2.156
	(1.111)	(1.322)
$k = +5$	-2.426***	-1.262
	(0.705)	(1.121)
$k = +6$	-1.874	-2.701***
	(1.230)	(0.898)
$k = +7$	-3.651***	-2.206**
	(1.181)	(0.800)
$k = +8$	-4.811***	-2.926***
	(1.057)	(0.801)
$k = +9$	-2.628***	-2.861***
	(0.608)	(0.468)
$k = +10$	-2.482**	-2.550***
	(0.946)	(0.713)
Observations	715,142	889,186
R^2	0.011	0.008

Notes: Interaction coefficients δ_k from equation (2) estimated separately by gender. Dependent variable: weekly effective hours among workers employed in services. Coefficients capture the differential response of high agricultural dependence regions relative to low dependence regions at each event time. HighAg is defined at the regional level (median split of 2010–2012 agricultural employment shares across the sixteen regions). Standard errors clustered at the region level (16 clusters) in parentheses; wild cluster bootstrap p-values reported in the text for the post-onset coefficients. Omitted period: $k = -1$. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

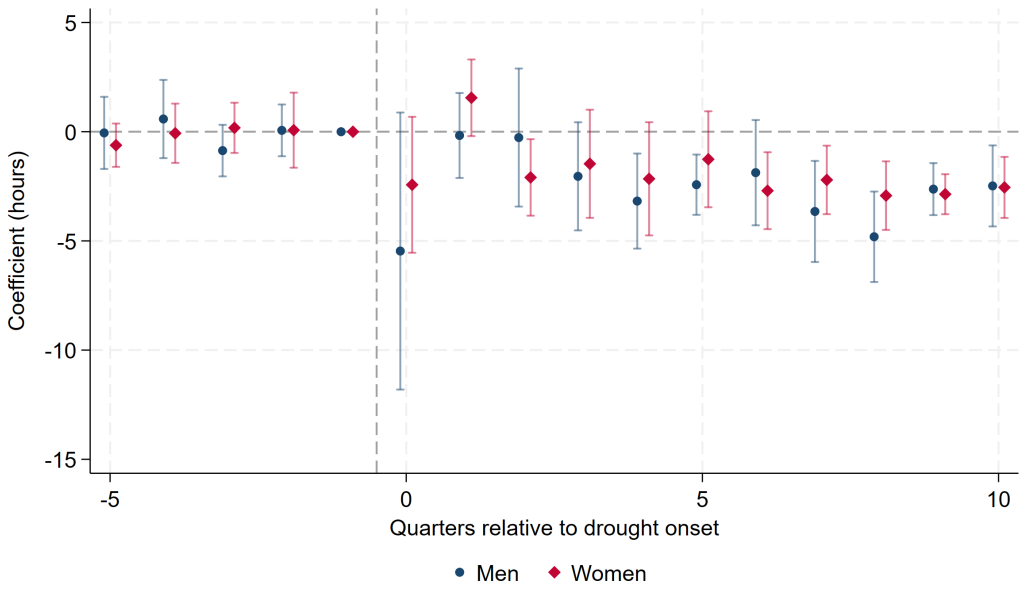


Figure 9: Spillover estimates: Effective hours in services (× high agricultural dependence)

Notes: Interaction coefficients δ_k from equation (2). Bars show 95% confidence intervals based on cluster robust standard errors (16 clusters). Omitted period: $k = -1$.

services should respond more strongly. Public and social services (education, health, public administration) provide a useful benchmark, as their funding is largely independent of local agricultural conditions, so significant effects in this branch would suggest a confounding force rather than the hypothesised channel.

The branch level results are broadly consistent with a local demand mechanism. Consumer market services (commerce and hospitality) show declines for women at several post-onset horizons, a pattern that aligns with women's disproportionate concentration in retail and hospitality employment. When drought reduces farm income and household expenditure, demand for these locally consumed services contracts, and women, who constitute the majority of workers in these branches, bear the adjustment. Transport and communications display noisier responses, reflecting the heterogeneous nature of this branch, in which some activities such as freight transport of agricultural products are directly exposed to farm output while others such as telecommunications are largely independent. Producer and business services show some significant movements, particularly for women at medium horizons, consistent with production linkage effects passing through professional services connected to agricultural trade and export logistics.

The most informative contrast is with public and social services, which display relative stability at short horizons and, at longer horizons ($k = +8$ to $+10$), positive differentials in more agriculture dependent regions. This pattern, diverging from the negative effects

Table 7: Service Branch Spillovers — Commerce and Hospitality

	Men	Women
$k = -5$	-2.725**	0.499
	(1.005)	(0.630)
$k = -4$	1.455	0.748
	(1.292)	(1.840)
$k = -3$	-0.197	0.281
	(1.233)	(2.313)
$k = -2$	-2.131	2.934
	(3.150)	(2.543)
$k = 0$	1.500	0.435
	(1.350)	(2.134)
$k = +1$	3.003**	1.170
	(1.129)	(1.217)
$k = +2$	0.991	2.464
	(2.336)	(2.443)
$k = +3$	0.479	0.898
	(1.837)	(2.752)
$k = +4$	4.968	-0.696
	(3.021)	(2.959)
$k = +5$	0.314	0.869
	(1.622)	(3.144)
$k = +6$	-4.614***	2.310
	(0.975)	(1.326)
$k = +7$	-4.347***	0.00792
	(0.887)	(1.169)
$k = +8$	7.054***	0.615
	(0.627)	(0.937)
$k = +9$	4.762***	-3.293*
	(0.831)	(1.603)
$k = +10$	1.997***	-2.283**
	(0.541)	(0.967)
Observations	129,952	152,332
R^2	0.066	0.038

Notes: Interaction coefficients δ_k from equation (2) estimated separately by gender. Dependent variable: weekly effective hours among workers employed in commerce and hospitality. HighAg defined at the regional level. Standard errors clustered at the region level (16 clusters) in parentheses. Omitted period: $k = -1$. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 8: Service Branch Spillovers — Transport and Communications

	Men	Women
$k = -5$	2.707*	-0.542
	(1.448)	(2.586)
$k = -4$	0.680	-5.383
	(1.007)	(6.298)
$k = -3$	-3.696**	3.895
	(1.333)	(2.673)
$k = -2$	-4.933**	6.486**
	(2.076)	(2.695)
$k = 0$	-1.150	-10.04**
	(1.949)	(4.248)
$k = +1$	-3.352	-9.999***
	(1.983)	(2.738)
$k = +2$	0.707	3.714
	(1.628)	(3.362)
$k = +3$	3.690	-1.584
	(4.276)	(2.602)
$k = +4$	5.433	-3.756
	(3.806)	(4.674)
$k = +5$	-0.330	0.00496
	(1.215)	(5.016)
$k = +6$	-1.164	12.47***
	(2.345)	(3.174)
$k = +7$	-0.264	0.688
	(1.475)	(3.449)
$k = +8$	7.218***	7.398***
	(0.910)	(2.238)
$k = +9$	-5.733***	-10.81***
	(0.688)	(2.535)
$k = +10$	-11.31***	-4.462***
	(0.624)	(1.326)
Observations	69,792	15,526
R^2	0.024	0.064

Notes: Interaction coefficients δ_k from equation (2) estimated separately by gender. Dependent variable: weekly effective hours among workers employed in transport and communications. HighAg defined at the regional level. Standard errors clustered at the region level (16 clusters) in parentheses. Omitted period: $k = -1$. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 9: Service Branch Spillovers — Finance and Business Services

	Men	Women
$k = -5$	0.604	-3.612**
	(2.081)	(1.648)
$k = -4$	0.679	-7.271***
	(2.206)	(2.409)
$k = -3$	-3.104*	0.0225
	(1.733)	(1.547)
$k = -2$	0.910	-1.970
	(2.533)	(2.013)
$k = 0$	-7.448*	4.130
	(3.823)	(4.862)
$k = +1$	-0.337	1.861
	(1.858)	(2.958)
$k = +2$	3.297**	5.680*
	(1.462)	(2.696)
$k = +3$	5.713	3.572
	(3.742)	(2.105)
$k = +4$	-1.841	5.886*
	(3.704)	(3.040)
$k = +5$	3.271***	-2.903
	(0.761)	(4.235)
$k = +6$	9.375**	-2.366
	(3.566)	(3.771)
$k = +7$	0.121	-1.841
	(2.727)	(3.293)
$k = +8$	-0.884	15.03***
	(1.555)	(1.315)
$k = +9$	4.284***	-0.436
	(0.885)	(2.064)
$k = +10$	3.585**	0.146
	(1.379)	(1.377)
Observations	43,921	38,533
R^2	0.039	0.024

Notes: Interaction coefficients δ_k from equation (2) estimated separately by gender. Dependent variable: weekly effective hours among workers employed in finance, real estate, and professional services. HighAg defined at the regional level. Standard errors clustered at the region level (16 clusters) in parentheses. Omitted period: $k = -1$. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 10: Service Branch Spillovers — Public and Social Services

	Men	Women
$k = -5$	-0.0868	0.852
	(1.144)	(0.806)
$k = -4$	1.959*	-1.451*
	(0.929)	(0.704)
$k = -3$	-0.916	-1.080
	(0.959)	(0.840)
$k = -2$	1.029	0.452
	(0.726)	(0.795)
$k = 0$	2.616	-1.857**
	(1.954)	(0.835)
$k = +1$	1.603	-0.0156
	(1.690)	(0.886)
$k = +2$	-1.385	2.664**
	(0.930)	(0.936)
$k = +3$	2.438*	1.296
	(1.215)	(1.259)
$k = +4$	2.844***	-0.107
	(0.919)	(1.103)
$k = +5$	-0.824	1.242
	(1.526)	(0.850)
$k = +6$	2.064	1.536
	(2.070)	(0.900)
$k = +7$	1.527	-0.167
	(1.334)	(0.863)
$k = +8$	1.030*	0.937***
	(0.506)	(0.275)
$k = +9$	4.879***	2.423***
	(0.890)	(0.329)
$k = +10$	6.312***	3.805***
	(0.510)	(0.262)
Observations	112,547	230,997
R^2	0.037	0.021

Notes: Interaction coefficients δ_k from equation (2) estimated separately by gender. Dependent variable: weekly effective hours among workers employed in education, health, and public administration. HighAg defined at the regional level. Standard errors clustered at the region level (16 clusters) in parentheses. Omitted period: $k = -1$. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

in private sector branches, is consistent with public employment acting as a partial stabiliser. When private demand contracts, public sector hours remain constant or expand through countercyclical fiscal responses, partially cushioning the aggregate labour market impact. The contrast between private and public branches provides indirect evidence that the spillover mechanism operates through market mediated demand and production linkages rather than through correlated shocks to all economic activity.

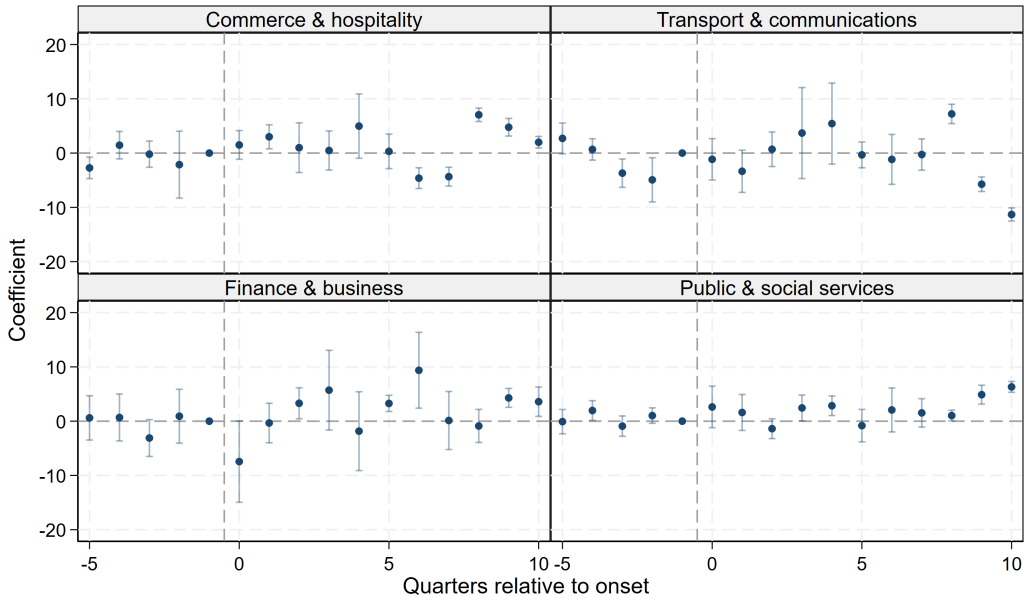


Figure 10: Service branch spillovers: Men

Notes: Interaction coefficients δ_k by service branch from equation (2), men only. Bars show 95% confidence intervals based on cluster robust standard errors (16 clusters). Omitted period: $k = -1$.

Two interpretive cautions apply to this branch level analysis. First, given the 16 regional clusters, individual event time coefficients in the branch level regressions are imprecisely estimated, and the branch level decomposition should be read as suggestive evidence about channels rather than as precise point estimates. The block level summaries in Section 7 provide more reliable aggregate inference. Second, the outcome is average hours among workers employed in each branch, so the estimates conflate within job hour adjustment with selection into and out of branch level employment.

6.4 Migration as a residual margin

The preceding results show that drought reduces labour input within agriculture and in connected non-agricultural sectors. A natural question is whether spatial reallocation, in particular internal migration, provides an additional adjustment margin, and whether the

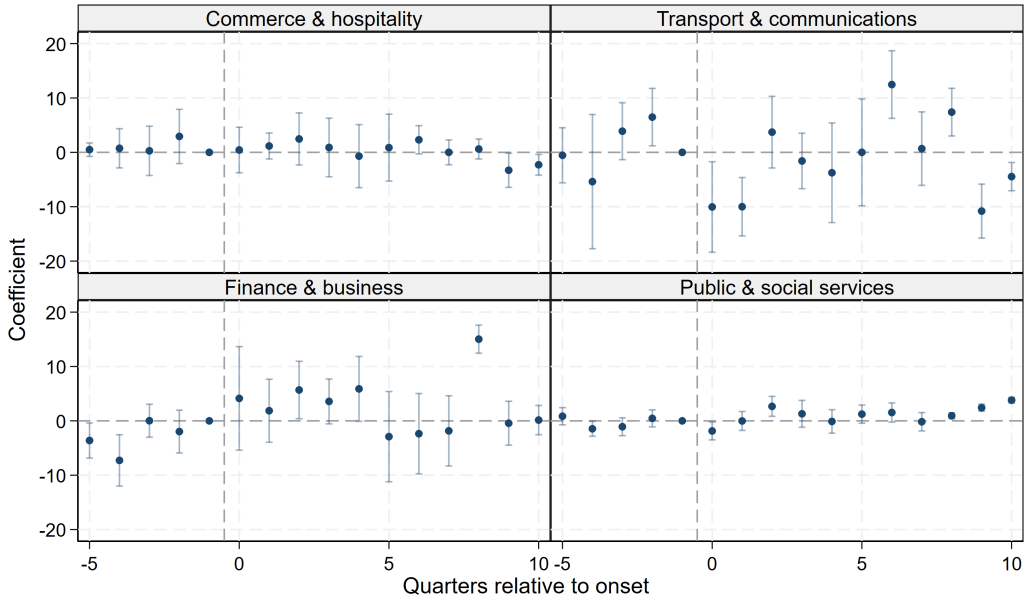


Figure 11: Service branch spillovers: Women

Notes: Interaction coefficients δ_k by service branch from equation (2), women only. Bars show 95% confidence intervals based on cluster robust standard errors (16 clusters). Omitted period: $k = -1$.

local economic structure shapes this response. Drought exposure at the origin is associated with a 2.1 percentage point increase in the probability of interregional migration ($p < 0.05$; Table 11). Given a baseline migration rate of approximately 3–4%, this effect represents a roughly 50–70% increase in the likelihood of moving across regions, a substantial response in proportional terms even though migration remains a relatively uncommon event in levels. Importantly, the drought migration relationship is attenuated in more diversified origins. The interaction coefficient (-0.035, $p < 0.05$) implies that a one standard deviation increase in the diversification index offsets most of the drought induced migration effect. This is consistent with diversified local economies offering reallocation opportunities that reduce the incentive to migrate, a mechanism emphasised by Jessoe et al. (2018) and Blakeslee et al. (2020) in other contexts. The gender interacted specification reveals that the drought migration relationship does not differ significantly between men and women, suggesting that migration reflects household relocation rather than gender specific labour migration.

In Table 12, I find no evidence that out migration buffers women's labour income. The interaction between the out migration rate and a female indicator is negative and significant (-1.40, $p < 0.05$), indicating that women's earnings move worse relative to men's as out migration rises. In economic terms, this coefficient implies that a 10 percentage point increase in the out migration rate is associated with a 14% decline in women's earnings relative to men's, a large differential. This finding is consistent with persistent occupational segmenta-

Table 11: Migration and Local Diversification

	(1) Baseline	(2) Gender interactions
Origin drought (SPEI-3 ≤ -1)	0.021^{**}	0.022^{**}
	(0.007)	(0.008)
Diversification at origin (1−HHI)	0.138^{***}	0.147^{***}
	(0.041)	(0.041)
Drought × Diversification	-0.035^{**}	-0.037^{*}
	(0.015)	(0.019)
Female	−0.001	0.007^{*}
	(0.001)	(0.003)
Female × Drought		−0.003
		(0.008)
Female × Diversification		-0.018^{*}
		(0.009)
Female × Drought × Diversification		0.004
		(0.019)
Observations	347,781	347,781
R^2	0.012	0.012

Notes: Linear probability estimates from equation (3). Dependent variable: indicator equal to one if the individual's current region differs from the region reported five years earlier. Controls: age (quadratic) and education dummies. Fixed effects: origin region and year. Standard errors clustered at the origin region in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

tion, as women's concentration in consumer facing services and seasonal agricultural tasks means that the demand contraction from drought is not offset by out migration, because the workers who leave are predominantly men in sectors where local demand is already stronger.

Table 12: Earnings and Out-Migration Among Stayers

	(1) Overall out-migration	(2) Male out-migration
Out-migration rate	0.179	0.578
	(0.535)	(0.460)
Female	-0.306^{***}	-0.309^{***}
	(0.034)	(0.034)
Out-migration rate × Female	-1.402^{**}	-1.284^{*}
	(0.653)	(0.616)
Observations	151,916	151,916
R^2	0.298	0.298

Notes: Estimates from the stayer earnings specification in equation (4). Dependent variable: log hourly wage proxy from the ESI module. Controls: age (quadratic) and education dummies. Fixed effects: origin region and year. Sample restricted to employed stayers whose current region of residence equals their origin region. Standard errors clustered at the origin region in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Two robustness checks reinforce this interpretation and address the concern that the headline coefficient relies on a single regression on a relatively small sample. First, I split the stayer sample by current sector of employment. The interaction between out migration and the female indicator is concentrated among stayers employed in services, where most working women in non-agriculture are observed. In the services subsample of roughly 96,000 observations, the interaction coefficient is -0.99 ($p < 0.05$), close in magnitude to the pooled estimate, while the agriculture and manufacturing subsamples have the same negative sign but coefficients that are imprecisely estimated given their smaller sizes. This pattern is consistent with the channel emphasised throughout the paper, namely that drought induced contractions in locally oriented services penalise women who stay because they are concentrated in those branches and have fewer alternative within region transitions. Second, I implement a temporal falsification by replacing contemporaneous out migration with the value lagged three years; the lagged interaction coefficient is similar in magnitude (-1.08) but only marginally significant given the autocorrelation of out migration rates across years and the loss of observations at the start of the panel, so the lagged placebo is informative but not decisive. Both robustness exercises are reported in Table 13.

Finally, Table 14 examines where migrants end up. Migrants are 1.4 percentage points more likely to be employed outside agriculture than observationally similar non-migrants

Table 13: Robustness: Stayer Earnings and Out-Migration — Sectoral Heterogeneity and Temporal Placebo

	Sample of stayer-earner observations				Placebo
	Baseline	Ag	Mfg	Services	Lag-3
Out-migration rate	0.190	−1.420	−0.150	−0.066	
	(0.563)	(1.505)	(0.964)	(0.497)	
Out-mig × Female	−1.598**	−0.200	−0.660	−0.990**	
	(0.674)	(1.276)	(1.775)	(0.432)	
Out-migration rate, lag −3					−0.518
					(0.894)
Out-mig (lag −3) × Female					−1.081
					(1.018)
Female	−0.293***	−0.282***	−0.448***	−0.284***	−0.322***
	(0.035)	(0.052)	(0.112)	(0.021)	(0.046)
Observations	141,743	13,158	32,607	95,639	60,818
R^2	0.252	0.199	0.218	0.263	0.239

Notes: Columns (1)–(4) re-estimate the stayer earnings specification of equation (4) on subsamples by current sector of employment. Column (5) is a temporal falsification that replaces contemporaneous out-migration with the lagged value three years earlier. Controls (age, age squared, education dummies) and origin region × year fixed effects included throughout. Standard errors clustered at the origin region level (16 clusters) in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

in the same destination region ($p < 0.05$). This pattern reinforces the interpretation that drought induced migration is part of a broader sectoral reallocation. Workers leave agriculture dependent origins and are absorbed into non-agricultural employment at their destinations, rather than into a mechanism that preserves agricultural employment structures. The absence of a significant migrant $\times$ female interaction suggests that the sectoral reallocation at destination does not differ meaningfully by gender, although the small migrant subsample limits the power to detect gender specific patterns.

Table 14: Destination Absorption into Non-Agricultural Employment

	(1) All employed	(2) Migrants only
Migrant (across regions, 5yr)	0.014**	
	(0.006)	
Female	0.042**	0.037***
	(0.014)	(0.009)
Migrant $\times$ Female	−0.005	
	(0.008)	
Origin drought (SPEI-3 ≤ -1)		0.006
		(0.010)
Destination diversification (1−HHI)		−0.129
		(0.378)
Observations	170,764	6,905
R^2	0.128	0.087

Notes: Linear probability estimates of an indicator for non-agricultural employment at destination. Fixed effects: destination region and year. Column (2) restricts the sample to employed migrants and adds origin region drought exposure and destination diversification as covariates. Standard errors clustered at the destination region in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

7 Robustness

Pre-treatment balance and placebo tests support the validity of the stacked design. Table 15 compares pre-treatment means of key covariates (gender composition, age, education, employment rates, and sectoral shares) between treated and clean control regions in the pre-event period. The two groups are balanced on observable characteristics, supporting the parallel trends assumption underlying the event study design. As a falsification test, I estimate equation (1) using public sector employment as the outcome (Table 16; Figure 12). Given that government employment is funded through national fiscal transfers rather than

local agricultural revenues, drought should not affect it through the hypothesised channel. For men, the placebo event time coefficients are small and statistically insignificant, with joint Wald p-values of 0.80 for the leads and 0.83 for the post-onset block, consistent with the absence of a drought channel for an outcome plausibly insulated from local agricultural conditions. For women, by contrast, several lead coefficients ($k = -4, -3, -2$) are positive and significant at the 5% level, raising the concern that the design may pick up correlated shocks for female outcomes.

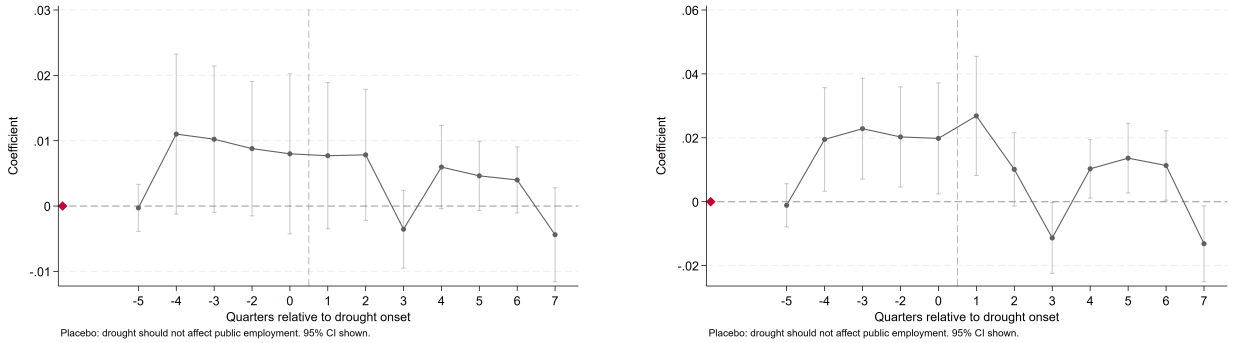
A natural reading of the women's placebo pattern attributes it to a secular feminisation of the Chilean services sector during 2010–2024, driven by the 2009 Equal Pay Act, subsequent gender parity directives within central government, and the expansion of social programs that operate through female staffed roles in education and health. Given that severe drought onsets are concentrated in the second half of the sample, this trend correlates mechanically with cohort calendar time rather than with the agricultural channel under examination. To probe this interpretation, I run two additional placebos using outcomes that are plausibly insulated from local agricultural demand and production linkages, namely employment in finance and business services (Santiago concentrated, serving national rather than local markets, and the most male dominated of the major services branches) and employment in transport and communications. Table 17 summarises the joint pre trends and post-onset p-values across all three outcomes. The pattern that emerges is informative. Men pass cleanly across all three placebos (pre trends $p = 0.80, 0.19, 0.30$; post-onset $p = 0.83, 0.17, 0.15$), so the male estimates in Sections 6.1–6.3 are not driven by correlated shocks. For women, pre trends remain significant for public and transport employment, consistent with feminisation operating across multiple services branches, but the cleanest of the three placebos, finance and business services, shows insignificant post-onset coefficients ($p = 0.24$), indicating that women's outcomes after drought onset are not contaminated by the same secular trend in branches that are less affected by it. Read together, the placebo evidence supports the male estimates unconditionally and the female estimates conditional on the recognition that pre trends in services branches reflect a real but unrelated secular trend rather than a drought confounder.

The findings are robust to alternative definitions of drought onset. I re-estimate the event study models under three variants: GAP8, which requires a minimum gap of eight quarters between consecutive onsets within a region and so reduces event overlap; TAU2, which uses a more extreme threshold of SPEI-3 ≤ -2.0 to select rarer and more severe episodes; and PERSIST2, which requires drought to persist for at least two consecutive quarters and so selects sustained episodes. Table A.2 in the Appendix reports the blocked estimates for the agricultural employment share. The main pattern, an immediate decline in women's

Table 15: Pre-Treatment Covariate Balance

Variable	Control	Treated	Difference	SE	p-value
Female (share)	0.498	0.499	−0.001	0.002	0.649
Age (mean, years)	38.205	38.298	−0.094	0.196	0.633
Secondary education	0.496	0.483	0.013	0.008	0.091
Tertiary education	0.314	0.310	0.004	0.014	0.749
Employed	0.626	0.630	−0.004	0.009	0.662
Employed in agriculture	0.100	0.115	−0.016	0.014	0.267
Employed in manufacturing	0.249	0.232	0.017	0.011	0.131
Employed in services	0.651	0.653	−0.002	0.013	0.893
Effective weekly hours	39.190	38.607	0.583	0.380	0.126
Observations (region-quarters)			255		

Notes: Pre-treatment means of selected covariates in treated and clean control regions over the event window. The "Difference" column reports the treated minus control difference, "SE" its standard error, and p-value the result of a t-test of equality of means clustered at the region level. Reported at the regional level for comparability with the spillover specification; the provincial-level balance produces an analogous pattern.



(a) Men **(b)** Women

Figure 12: Placebo event study: Public sector employment

Notes: Event time coefficients β_k from equation (1) with public sector employment as the outcome, estimated separately by gender. Bars show 95% confidence intervals based on cluster robust standard errors (54 provincial clusters). Omitted period: $k = -1$.

Table 16: Placebo Test — Public Sector Employment

	Men	Women
$k = -5$	0.000332	-0.000489
	(0.00175)	(0.00261)
$k = -4$	0.00650	0.0110**
	(0.00509)	(0.00457)
$k = -3$	0.00576	0.0138***
	(0.00482)	(0.00432)
$k = -2$	0.00548	0.0129***
	(0.00452)	(0.00438)
$k = 0$	0.00467	0.00958*
	(0.00468)	(0.00494)
$k = +1$	0.00400	0.0176***
	(0.00419)	(0.00536)
$k = +2$	0.00618	0.00925*
	(0.00457)	(0.00527)
$k = +3$	0.000481	-0.00268
	(0.00274)	(0.00349)
$k = +4$	0.00521	0.00918**
	(0.00380)	(0.00407)
$k = +5$	0.00410	0.0128***
	(0.00327)	(0.00456)
$k = +6$	0.00392	0.0117**
	(0.00333)	(0.00478)
$k = +7$	0.000181	-0.00342
	(0.00287)	(0.00323)
Observations	15,129,631	11,969,185
R^2	0.148	0.353

Notes: Coefficients β_k from equation (1) estimated separately by gender. Dependent variable: indicator for employment in public sector services (education, health, public administration). Reported event-time horizons are restricted to those with non-degenerate variation in the placebo sample. Standard errors clustered at the province level (54 clusters) in parentheses. Omitted period: $k = -1$. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 17: Triple Placebo — Block Wald Tests Across Three Insulated Outcomes

	Men		Women	
Outcome	Pre-trends p	Post joint p	Pre-trends p	Post joint p
Public sector employment	0.799	0.826	0.027**	0.036**
Finance & business services	0.187	0.170	0.097*	0.235
Transport & communications	0.303	0.145	0.011**	0.060*

Notes: Joint Wald p-values from event-study placebo regressions of equation (1) estimated separately by gender on three outcomes plausibly insulated from the agricultural drought channel: public sector employment (education, health, public administration), finance and business services (Santiago concentrated, serves national markets), and transport and communications. Pre-trends test joint significance of $\beta_{-5}, \beta_{-4}, \beta_{-3}, \beta_{-2}$; post joint tests joint significance of $\beta_0, \beta_{+1}, \ldots, \beta_{+10}$. Standard errors clustered at the province level (54 clusters). * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

agricultural employment share and a delayed decline for men, is robust across specifications. Under the gap8 definition, women's employment share declines by 1.3 percentage points at horizons 7–10 ($p < 0.05$). The persist2 definition yields a sharper immediate response for women (-3.4 pp at $k = 0$ to $+2$, $p < 0.01$), consistent with sustained drought episodes producing larger effects. Pre trends are clean under gap8 and the baseline but noisier under tau2, reflecting the smaller number of extreme events.

To address the small number of clusters, I summarise the event study dynamics into post-onset blocks of short run (quarters 0–2), medium run (quarters 3–6), and longer run (quarters 7–10). For each block, I report the average coefficient, its clustered standard error, and a Wald test for joint significance of the constituent event time coefficients, and complement these with wild cluster bootstrap (WCB) p-values (Roodman et al., 2019) computed using 9,999 replications and a fixed seed for reproducibility. The block level results confirm the spillover findings, since medium- and longer run block averages for services hours are negative and significant for both genders in agriculture dependent regions, and WCB p-values do not change these conclusions at conventional significance levels. For the contract composition mechanism (Table A.3), the regional non-stacked block average specification yields a medium run rise in men's limited term contract share of around two to three percentage points, with significance varying across alternative onset definitions ($p < 0.01$ under the baseline, $p < 0.05$ under gap8, not significant under the more demanding persist2 definition). In the provincial stacked main specification the analogous medium run block average for men is similar in sign but smaller in magnitude and not statistically significant, indicating that the contract margin moves in the direction predicted by the model but its precise magnitude depends on the cohort and fixed effect structure used to identify it.

The spillover specification compares regions above and below the median agricultural

employment share, which raises the concern that the eight versus eight regional comparison may capture other structural differences across regions correlated with agricultural dependence. I address this concern with three alternative specifications reported in Table 18. The first replaces the binary indicator with the continuous agricultural share base variable, exploiting all the regional variation rather than half against half. The negative interaction effects in services hours persist with magnitudes consistent with the binary baseline, and pre trends remain clean for men ($p = 0.29$) and only marginally suggestive for women ($p = 0.08$). The second adds interactions of event time with regional baseline characteristics that could plausibly correlate with agricultural dependence, namely manufacturing share, tertiary education share, and the employment rate; the post-onset interaction coefficients on agricultural share base retain the same sign and similar magnitudes for men, but pre trends deteriorate, reflecting the cost of saturating the model with controls when only sixteen regions are available, and I report this specification as a stringent stress test rather than as a preferred specification. The third defines high and low agricultural dependence as the top and bottom terciles of the regional agricultural share distribution, dropping the middle tercile, and so compares structurally more distinct regions. The magnitudes are very close to the binary baseline, with services hours falling by roughly three to four hours per week in the high tercile relative to the low for men through ten quarters after onset, and pre trends are clean for both genders. Across the three specifications, the qualitative finding that services hours decline differentially in agriculture dependent regions is preserved, which suggests the headline result does not hinge on the median split.

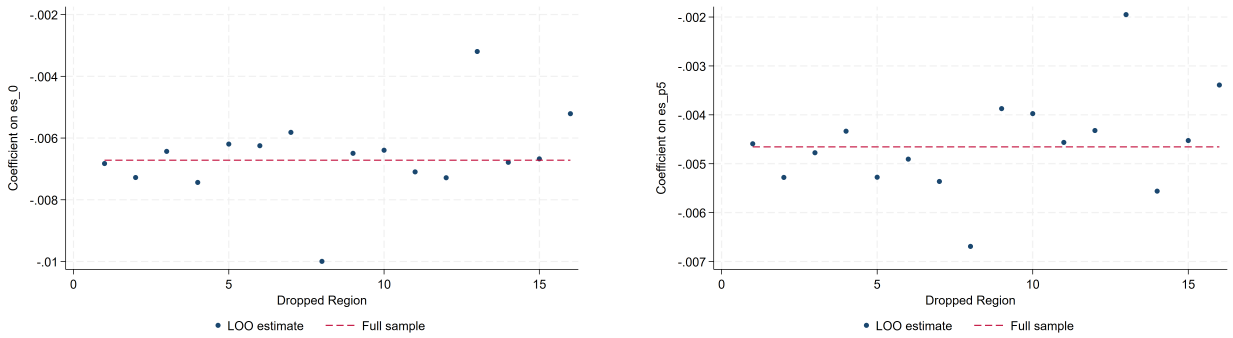
Two further checks address concerns about cluster sensitivity and within sample heterogeneity. Leave one out estimation, in which the regional version of the baseline specification for the agricultural employment share is re-estimated while sequentially dropping each of the 16 regions, produces stable point estimates with no single region exerting disproportionate influence on the impact and medium run effects (Figure 13). Heterogeneity by education (tertiary versus non-tertiary) and age group (15–30, 31–45, 46–65) is reported in Tables A.4 and A.5 in the Appendix. The agricultural employment response is concentrated among non-tertiary educated workers, consistent with the composition of seasonal agricultural labour, and among age groups younger workers (15–30) show larger extensive margin effects, reflecting their greater reliance on seasonal entry level contracts.

The migration and earnings results are also robust to alternative specifications. The migration estimates are robust to using SPEI-3 ≤ -1.5 instead of ≤ -1.0 as the drought measure, and to alternative constructions of the migration rate (region year averages versus individual level indicators). The earnings interaction (out migration $\times$ female) remains negative and significant across specifications, and the lack of wage buffering for women

Table 18: Robustness: Services Spillovers — Alternative HighAg Specifications

	Men		Women	
Specification / Block	Coef.	SE	Coef.	SE
Spec 1: Continuous HighAg (× ag share base)				
Pre-trends p (Wald)	0.294		0.083	
Post 0–2	-20.5^{***}	(6.69)	-10.2^{***}	(3.77)
Post 3–6	-19.7^{***}	(3.42)	-15.3^{***}	(5.02)
Post 7–10	-23.8^{***}	(2.82)	-18.4^{***}	(1.86)
Spec 2: Continuous + region controls × event-time				
Pre-trends p (Wald)	0.000		0.052	
Post 0–2	-15.8	(31.7)	-14.1	(16.6)
Post 3–6	-25.8^{**}	(11.7)	-36.1^{**}	(17.5)
Post 7–10	-23.4^{**}	(11.4)	-4.9	(17.8)
Spec 3: Tercile (top vs bottom 1/3, dropping middle)				
Pre-trends p (Wald)	0.095		0.538	
Post 0–2	-3.08^{*}	(1.62)	-1.36^{*}	(0.79)
Post 3–6	-3.02^{***}	(1.03)	-1.96	(1.32)
Post 7–10	-3.89^{***}	(0.99)	-2.68^{***}	(0.83)

Notes: Block averages of post-onset interaction coefficients in services hours specification (equation 2), comparing three definitions of regional agricultural dependence. Spec 1 replaces the binary *HighAg* with continuous *ag share base* (2010–2012); coefficients are interpreted as the marginal effect per unit of ag share, so a 10 percentage point difference corresponds to roughly one tenth of the reported coefficient. Spec 2 adds interactions of event time with regional baseline manufacturing share, tertiary education share, and employment rate. Spec 3 compares regions in the top tercile of *ag share base* to those in the bottom tercile, dropping the middle tercile. Pre-trends p-values are joint Wald tests on $k = -5, -4, -3, -2$ interaction coefficients. Standard errors clustered at the region level (16 clusters). * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.



(a) Impact effect ($k = 0$)

(b) Medium run effect ($k = +5$)

Figure 13: Leave one out estimates: Agricultural employment share

Notes: Each marker shows the baseline event study coefficient for the agricultural employment share when the indicated region is dropped from the sample. Vertical bars are 95% confidence intervals based on cluster robust standard errors. The horizontal line marks the full sample point estimate.

is not sensitive to the inclusion of additional controls (sector fixed effects, urban or rural indicators) or to restricting the sample to full time workers.

8 Conclusion

This paper estimates how severe drought onsets reshape labour market outcomes and gender gaps in Chile over 2010–2024. I link quarterly labour force survey microdata to region level drought exposure and implement a stacked event study design with multiple onsets, providing joint evidence on direct agricultural adjustment, spillovers to non-agricultural sectors, and spatial reallocation.

Three findings emerge. First, within agriculture, hours among continuing workers change little at onset, but women's agricultural employment share falls on impact and men experience a delayed shift toward less stable contracts. The stability of hours among survivors, combined with sharp extensive margin responses, points to selection, as the seasonal matches that would otherwise be created or renewed at the margin are the first to disappear, leaving average hours and wages among workers largely unchanged. Second, drought propagates to the broader regional economy. Services, and to a lesser extent manufacturing, exhibit persistent reductions in hours in regions more dependent on agriculture, and the within services decomposition suggests that local demand contractions in consumer facing branches and production linkage effects in transport and business services both play a role, while public provision partially stabilises labour input. Third, migration acts as a residual margin of adjustment that is weaker where origin regions are more diversified, and higher out migration does not translate into improved earnings for women who remain, consistent with continued

weak demand in the segments where women are concentrated.

These findings carry direct implications for policy. Drought response strategies that focus exclusively on agricultural production risk missing important spillovers to services and to job quality. Protecting women's attachment to paid work during climatic stress may require targeted instruments beyond farm support, including childcare provision, transport subsidies, and active labour market policies that facilitate transitions into stable non-agricultural employment. Strengthening local public services and social protection can also help stabilise demand and reduce distress migration in less diversified regions, particularly where the contraction is concentrated in consumer facing services on which women rely for paid work.

Several limitations should be acknowledged. First, while the direct effects specifications use 54 provincial clusters, the spillover specifications still rely on the sixteen regional clusters within which baseline agricultural dependence varies, which limits the power of individual hypothesis tests for those coefficients despite the large sample of individual observations. Second, the intensive margin estimates condition on continued employment, so they are subject to survivor bias, and if drought selectively removes workers with lower or higher hours, average hours among survivors may not reflect the causal effect on a fixed population. Third, migration is measured retrospectively over a five year window, which limits the ability to link specific drought events to migration decisions at high frequency. Fourth, the spillover identification relies on cross regional variation in baseline agricultural dependence, which could correlate with other structural differences across regions, and although the fixed effects structure absorbs level differences, differential trends remain a potential concern.

Two directions for future work follow from these limitations. First, complementary microdata on firm level adjustments, hiring decisions, contract offers, and hours rationing would help disentangle demand side from supply side responses to drought within services, a distinction that the current analysis can only approach indirectly through the within services decomposition. If consumer facing firms reduce hiring because local spending falls while transport and logistics firms cut hours because agricultural volumes decline, the policy responses differ, and firm level data would identify which mechanism dominates. Second, extending the analysis to exploit finer spatial variation at the municipality level, where economic mechanisms could plausibly differ across municipalities within a province, would provide a more demanding test of the local demand and production linkage channels, and would help assess whether the attenuated migration response in diversified regions reflects genuine within region reallocation or simply lower exposure to the initial shock.

Appendix

A.1 Formal framework

Consider a small open regional economy with a continuum of workers of mass one, indexed by gender $g \in \{m, f\}$ with population shares μ_m and $\mu_f = 1 - \mu_m$. Production takes place in two sectors, agriculture (A) and non-agriculture (N). Output in agriculture is given by $Y_A = z \cdot L_A^{\alpha}$ with $\alpha \in (0, 1)$, where z is total factor productivity directly exposed to climate. Output in non-agriculture is $Y_N = q(z; \theta) \cdot L_N^{\beta}$ with $\beta \in (0, 1)$, where the term $q(z; \theta)$ captures the dependence of local non-agricultural activity on agricultural conditions, formalised as

$$q(z; \theta) = \bar{q}_0 + \theta \cdot z^{\eta}, \qquad \eta > 0.$$

Here $\bar{q}_0 > 0$ is autonomous demand for non-agricultural output that is independent of local agricultural conditions, while $\theta \in [0, \bar{\theta}]$ indexes the strength of the local linkage between agriculture and the rest of the regional economy through the income and production channels described in Section 3. Regions with higher θ correspond to higher baseline agricultural dependence. Labour markets clear competitively in each sector.

Agricultural employment is organised through two contract types. A fraction λ_g of workers of gender g employed in agriculture hold limited term contracts that expire automatically at the end of the production cycle, while the remaining $1 - \lambda_g$ hold open ended contracts protected by a fixed severance cost $s > 0$ in the event of dismissal. The institutional and descriptive evidence in Sections 2–4 motivates the asymmetry $\lambda_f > \lambda_m$, since women are disproportionately concentrated in seasonal arrangements that align with the harvest cycle. Workers in non-agriculture are not subject to this margin in the model since the seasonal contract structure is concentrated in agriculture.

In the unshocked benchmark, profit maximisation by representative firms yields agricultural and non-agricultural employment levels

$$L_A^* = \left(\frac{\alpha z}{w_A} \right)^{\frac{1}{1-\alpha}}, \qquad L_N^* = \left(\frac{\beta \, q(z; \theta)}{w_N} \right)^{\frac{1}{1-\beta}},$$

where w_A and w_N are the equilibrium wages, determined by labour market clearing given the gender specific outside options $\bar{u}_g$. In what follows I treat wages as approximately fixed in the relevant range, an assumption that is consistent with the empirical finding that wages and hours among continuing workers respond little to drought (Section 6.1) and that simplifies the comparative statics without affecting their qualitative content.

A severe drought is modelled as a fall in agricultural productivity from z_0 to $z_1 = (1-\delta)z_0$

with $\delta \in (0,1)$. Two adjustment margins follow.

In agriculture, a fall in z reduces the marginal product of labour at every employment level, so firms wish to contract employment. The contract structure determines how this contraction is distributed. Limited term workers can be released at zero cost simply by not renewing contracts; open ended workers can only be released at cost s per worker. For a sufficiently small shock δ, all adjustment occurs through non-renewal of limited term contracts and no open ended worker is dismissed. By gender, the change in agricultural employment is

$$\frac{dL_A^g}{dz} = \lambda_g \cdot \frac{dL_A^L}{dz} + (1-\lambda_g) \cdot \frac{dL_A^O}{dz},$$

where for small shocks $dL_A^O/dz \approx 0$. Since $dL_A^L/dz > 0$ (employment falls when productivity falls), and $\lambda_f > \lambda_m$, the gender specific responses satisfy $|dL_A^f/dz| > |dL_A^m/dz|$, with the magnitude of the gap increasing in $\lambda_f - \lambda_m$.

In non-agriculture, the productivity shock transmits indirectly through $q(z;\theta)$. Differentiating L_N^* with respect to z holding w_N fixed,

$$\frac{dL_N^*}{dz} = \frac{1}{1-\beta} \cdot \frac{L_N^*}{q(z;\theta)} \cdot \frac{\partial q}{\partial z} = \frac{1}{1-\beta} \cdot \frac{L_N^*}{q(z;\theta)} \cdot \theta\, \eta\, z^{\eta-1} > 0,$$

so a fall in z (drought) reduces non-agricultural employment, with the magnitude proportional to θ. Spillovers are mechanically larger in regions with higher baseline agricultural dependence. If women are more concentrated than men in the subset of non-agricultural branches more responsive to q (typically commerce and hospitality, where local demand is most relevant), then the gender specific spillover responses inherit the sign and magnitude of the underlying θ heterogeneity.

Workers may relocate to outside regions at a fixed cost c if the local utility net of cost falls below the outside option. Local utility is a function of expected local employment opportunities, summarised by the diversification index D, the breadth of employment available outside agriculture and outside the most exposed non-agricultural branches. Migration probability after the shock is

$$P_M(z;D) = \Phi\big(c - a\big[\Delta U^{loc}(z;D)\big]\big),$$

with Φ a smooth cumulative distribution of moving costs and $a > 0$ a scaling parameter. Higher D raises within region reallocation possibilities, partially offsetting the drought induced fall in local utility, so $\partial^2 P_M/\partial z\,\partial D < 0$, that is, drought induced out migration is attenuated in more diversified regions.

The model aggregates manufacturing and services into a single non-agricultural sector

N for tractability, while the empirical analysis in Sections 6.2–6.3 restores this distinction by estimating separate spillover specifications for manufacturing and for services and by further decomposing services into four branches. Within the framework above, this empirical decomposition can be read as exploiting heterogeneity in the elasticity η of local demand transmission across branches rather than introducing additional sectors with distinct production technologies. Branches more exposed to local consumer demand, in particular commerce and hospitality, correspond to higher η and so should display larger spillover responses; branches tied to upstream and downstream agricultural processing, including manufacturing and transport, correspond to intermediate η with effects that materialise more slowly as production volumes pass through the supply chain; branches serving national rather than local markets, such as finance and business services, correspond to lower η; and institutionally insulated branches, such as public services and the education and health subsectors, correspond to η close to zero. The two sector model therefore captures the average spillover effect that operates through the local demand and linkage channels.

Table A.2: Robustness: Agricultural Employment Share — Alternative Onset Definitions

	Men		Women	
Specification / Block	Coef.	SE	Coef.	SE
BASE ($\tau = -1.5$, $gap = 0$)				
Pre-trends p (Wald)	0.027		0.244	
Post 0–2	0.383	(0.567)	0.796	(1.163)
Post 3–6	−0.417	(0.574)	−0.668	(0.721)
Post 7–10	0.429	(0.507)	−1.457***	(0.309)
GAP8 (min gap = 8q)				
Pre-trends p (Wald)	0.175		0.541	
Post 0–2	0.353	(0.540)	0.509	(1.091)
Post 3–6	−0.587	(0.493)	−0.589	(0.831)
Post 7–10	0.238	(0.532)	−1.310**	(0.450)
PERSIST2 (drought persists $\geq 2q$)				
Pre-trends p (Wald)	0.171		0.000	
Post 0–2	−0.324	(1.005)	−3.357***	(0.948)
Post 3–6	−1.301*	(0.715)	−1.064	(2.519)
Post 7–10	0.382	(0.430)	−2.458***	(0.652)

Notes: Block averages of event-time coefficients from equation (1). Standard errors clustered at the province level (54 clusters). Pre-trends p-values from joint Wald test of all pre-onset coefficients. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A.3: Robustness: Limited-Term Contract Share — Alternative Onset Definitions

	Men		Women	
Specification / Block	Coef.	SE	Coef.	SE
BASE ($\tau = -1.5$, $gap = 0$)				
Pre-trends p (Wald)	0.370		0.415	
Post 0–2	0.695	(1.036)	0.030	(1.800)
Post 3–4	2.877***	(0.531)	2.169	(2.507)
GAP8 (*min gap* = 8q)				
Pre-trends p (Wald)	0.389		0.733	
Post 0–2	0.647	(1.106)	−0.686	(1.979)
Post 3–4	1.910**	(0.806)	1.048	(2.925)

Notes: Block averages of event-time coefficients from the non-stacked regional specification (cohort-pooled identifying variation, region-level fixed effects, 16 regional clusters). The provincial stacked specification used in the main mechanism table (Table 4) yields block averages of similar sign but smaller magnitude and weaker statistical significance, as discussed in the text. Event window is shorter for mechanism outcomes due to data availability; blocks 7–10 are not reported. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A.4: Heterogeneity by Education — Agricultural Employment Share

	Non-Tertiary		Tertiary	
	Men	Women	Men	Women
$k = -5$	0.00199	-0.00740*	-0.0109**	-0.00260*
	(0.00366)	(0.00409)	(0.00380)	(0.00134)
$k = -4$	0.0000602	-0.0123*	-0.0122**	-0.00234
	(0.00322)	(0.00599)	(0.00430)	(0.00282)
$k = -3$	0.0000873	-0.0106**	-0.0114**	-0.00972**
	(0.00500)	(0.00430)	(0.00452)	(0.00391)
$k = -2$	0.00445	-0.0131	-0.00903**	-0.00459
	(0.00453)	(0.00799)	(0.00349)	(0.00289)
$k = 0$	-0.00543	-0.0136*	-0.00914**	-0.00820***
	(0.00423)	(0.00679)	(0.00380)	(0.00264)
$k = +1$	-0.0111*	-0.0183**	-0.0126***	-0.00846*
	(0.00532)	(0.00679)	(0.00385)	(0.00414)
$k = +2$	-0.00304	-0.00852	-0.0103**	-0.00501
	(0.00659)	(0.00960)	(0.00375)	(0.00362)
$k = +3$	-0.00366	-0.00544	-0.00790	-0.00689**
	(0.00537)	(0.00541)	(0.00537)	(0.00318)
$k = +4$	-0.00513	-0.00134	-0.0134**	-0.00261
	(0.00619)	(0.00672)	(0.00461)	(0.00218)
$k = +5$	-0.0103	-0.00892	-0.0107**	-0.00320
	(0.00710)	(0.00636)	(0.00392)	(0.00339)
$k = +6$	-0.00273	-0.00304	-0.0111**	-0.00196
	(0.00943)	(0.0107)	(0.00450)	(0.00366)
$k = +7$	-0.00191	-0.00352	-0.00896*	-0.000890
	(0.00715)	(0.00540)	(0.00442)	(0.00154)
$k = +8$	-0.00242	-0.00388	-0.00777	-0.00574
	(0.00835)	(0.00617)	(0.00517)	(0.00343)
$k = +9$	-0.00180	-0.000965	-0.00630	-0.00178
	(0.0108)	(0.00857)	(0.00451)	(0.00328)
$k = +10$	-0.000689	0.000826	-0.00820**	-0.00387
	(0.00923)	(0.0124)	(0.00332)	(0.00388)
Observations	4,031,171	2,799,433	1,844,373	1,865,976
R^2	0.094	0.068	0.024	0.012

Notes: Coefficients β_k from equation (1) estimated separately by gender and education group. Dependent variable: indicator for employment in agriculture. Standard errors clustered at the region level (16 clusters) in parentheses. Omitted period: $k = -1$. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A.5: Heterogeneity by Age Group — Agricultural Employment Share

	15–30		31–45		46–65	
	Men	Women	Men	Women	Men	Women
$k = -5$	0.00736*	-0.00866*	-0.00125	-0.00577	-0.00555	-0.00386
	(0.00362)	(0.00473)	(0.00468)	(0.00366)	(0.00510)	(0.00347)
$k = -4$	-0.00242	-0.0106	-0.00174	-0.00633	-0.00866**	-0.00927**
	(0.00724)	(0.00683)	(0.00391)	(0.00444)	(0.00340)	(0.00406)
$k = -3$	-0.00440	-0.0135*	-0.00638	-0.0138**	-0.000271	-0.00548
	(0.00862)	(0.00657)	(0.00393)	(0.00508)	(0.00695)	(0.00321)
$k = -2$	0.000923	-0.00966	0.00374	-0.0112	-0.00403	-0.00773
	(0.00782)	(0.00697)	(0.00316)	(0.00734)	(0.00438)	(0.00568)
$k = 0$	-0.00982	-0.0146**	-0.00831	-0.00844*	-0.00565	-0.0116**
	(0.00723)	(0.00589)	(0.00529)	(0.00465)	(0.00383)	(0.00510)
$k = +1$	-0.0138*	-0.0197**	-0.00746	-0.0191**	-0.0118**	-0.00583
	(0.00697)	(0.00725)	(0.00587)	(0.00667)	(0.00404)	(0.00445)
$k = +2$	-0.00860	-0.0159*	-0.00589	-0.00747	-0.00154	0.0000358
	(0.00742)	(0.00873)	(0.00683)	(0.00760)	(0.00507)	(0.00736)
$k = +3$	-0.00594	-0.00968**	-0.00330	-0.00242	-0.00314	-0.00690
	(0.00615)	(0.00376)	(0.00735)	(0.00429)	(0.00577)	(0.00412)
$k = +4$	-0.00964	-0.0136**	-0.00385	0.00298	-0.00726	0.000777
	(0.00648)	(0.00561)	(0.00726)	(0.00739)	(0.00475)	(0.00278)
$k = +5$	-0.00411	-0.00982	-0.00866	-0.00589	-0.0114*	-0.00608
	(0.00860)	(0.00655)	(0.00597)	(0.00627)	(0.00608)	(0.00490)
$k = +6$	0.00205	-0.00922	-0.00266	-0.00320	-0.00953	0.000173
	(0.0125)	(0.00881)	(0.00881)	(0.00993)	(0.00641)	(0.00681)
$k = +7$	0.00878	-0.00653	-0.0112	-0.00557	-0.00234	0.00212
	(0.00590)	(0.00408)	(0.00766)	(0.00550)	(0.00589)	(0.00358)
$k = +8$	0.00491	-0.0142*	-0.00976	-0.00167	-0.00328	-0.00334
	(0.00604)	(0.00754)	(0.00826)	(0.00817)	(0.00620)	(0.00369)
$k = +9$	-0.00246	-0.0107	0.000794	-0.00184	-0.00552	0.00502
	(0.0110)	(0.00907)	(0.0107)	(0.00753)	(0.00806)	(0.00586)
$k = +10$	0.00117	-0.00699	-0.00181	-0.00416	-0.00394	0.00429
	(0.0128)	(0.00736)	(0.00973)	(0.0116)	(0.00757)	(0.00726)
Observations	1,392,010	1,084,830	1,946,382	1,686,531	2,537,152	1,894,048
R^2	0.103	0.069	0.139	0.082	0.168	0.073

Notes: Coefficients β_k from equation (1) estimated separately by gender and age group. Dependent variable: indicator for employment in agriculture. Standard errors clustered at the region level (16 clusters) in parentheses. Omitted period: $k = -1$. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

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