Two Margins of Labor Force Entry

Aditi Singh*
CAFRAL, Reserve Bank of India

Abstract

People outside the labor force can enter through employment or unemployment. We ask whether job entry reveals participation movements that unemployment-entry history misses. Using the Current Population Survey data for the US, we use principal component analysis to ascertain whether stronger movement into jobs is associated with subsequent entry into unemployment within this population. We find that relative employment entry among nonparticipants without a recorded yes to wanting a job is significantly associated with more unemployment entry twelve months later. After seasonal and dynamic controls, a one-standard-deviation increase corresponds to 0.096 percentage points, about 7.0% of mean entry, thus indicating that the two margins contain complementary information.

Keywords: Nonparticipation; employment transitions; labor-market attachment.
JEL classification: E24, J22, J64.

*We gratefully acknowledge the excellent support provided by my research assistant Yash Maurya. The views expressed here are solely the responsibility of the author and should not be interpreted as reflecting the views of the Centre for Advanced Financial Research and Learning (CAFRAL) or the Reserve Bank of India.

1 Introduction

Unemployment entry gives an incomplete view of how people outside the labor force connect with work. A nonparticipant observed in a job the following month is recorded entering through employment. A nonparticipant observed unemployed enters without a job. Both flows contribute to participation, but only the second adds to unemployment. This distinction matters for interpreting unemployment: entry into unemployment from nonparticipation adds to both the number unemployed and the labor force. Employment entry records the other route, in which participation comes with a job.

We ask whether employment entry contains information about later unemployment entry that the latter's own history misses. Our focus is the broad population of nonparticipants without a recorded yes to wanting a job. Their employment transitions reveal a connection to work that their current participation status does not record. Employment entry records a successful match with a job; unemployment entry records participation without one. Changes in job opportunities or personal circumstances may appear in these margins at different dates. The economic question is whether stronger movement into jobs is associated with subsequent entry into unemployment within this population.

Employment entry does contain such additional information in the observed sample. A one-standard-deviation increase in employment entry relative to movements shared with other nonemployed groups is associated with 0.096 percentage points more monthly unemployment entry twelve months later. This is about 7.0% of mean entry. The association survives calendar-month controls and a year of unemployment-entry history. The rates describe different monthly populations; they do not trace the same people from jobs into unemployment.

This question connects research on participation flows with evidence on differences in attachment to work. Elsby et al. [2015] establish the importance of participation flows for unemployment fluctuations. Hornstein et al. [2014] measure labor utilization using employment-transition probabilities, while Kudlyak and Lange [2018] show that labor-force histories predict job finding beyond reported desire for work. Graves and Huckfeldt [2026] compare job finding among active searchers and comparable nonparticipants. We examine what one observed route into the labor force tells us about the other. The contribution is evidence that unemployment-entry history alone does not summarize the subsequent participation movements of this broad group.

2 Data

We measure the two routes into the labor force using monthly US Current Population Survey records from January 1994–March 2026, obtained through IPUMS. The sample retains civilians aged sixteen or older with reported, rather than allocated, employment status. Consecutive-month matches use person identifiers with survey-rotation and demographic

checks [Rivera Drew et al., 2014]. Matched pairs receive the average of their two monthly final survey weights, restricted to positive, nonmissing values. Groups are defined in the first month and transitions dated by the second.

Five origin groups distinguish different connections to work among people without jobs. We separate unemployment lasting at least 27 weeks from shorter or unrecorded durations. Nonparticipants who want a job are divided by recorded search or no search during the preceding year; unknown search responses enter neither group. Our primary group, *other nonparticipants*, has no recorded yes to wanting a job. It includes explicit no answers, retirement or disability responses, and unanswered or skipped questions. Each entry rate divides the weighted count reaching its destination by the weighted total of matched origin-group members.

Two population checks ask whether the relationship extends to narrower definitions of nonparticipation. The explicit-no check retains only respondents explicitly saying they do not want a job in the primary group, leaving the other four groups unchanged. The prime-age check restricts all five groups to ages 25–54. Each check reconstructs the common employment-entry factor for its population.

3 Empirical strategy

We compare the primary group's job entry with movements shared across the five groups to distinguish its relative strength from their common movement. Standardizing each group's log employment-entry rate gives z_{gt}. Its first principal-component loading is λ_g, and $f_t = \sum_g \lambda_g z_{gt}$ summarizes the common movement. Relative employment entry is $\eta_t = z_{gt} - \lambda_g f_t$: a higher value means stronger entry than the group's loading on the common factor implies. Each population's standardization and principal components use its 379 complete, positive-rate months.

Employment entry adds useful descriptive information if its association survives the seasonal patterns and persistence already present in unemployment entry. Let q_t denote the primary group's monthly unemployment-entry rate. We estimate

$$q_{t+12} = \alpha + \beta\eta_t + \gamma f_t + \sum_{j=0}^{12} \rho_j q_{t-j} + \delta_{m(t)} + u_{t+12}. \tag{1}$$

The eleven calendar-month indicators $\delta_{m(t)}$ account for recurring seasonal movements. Current unemployment entry and twelve monthly lags summarize its recent path. Thus β asks whether employment entry distinguishes months with similar observed unemployment-entry histories, seasonal timing, and common employment movements. All leads and lags use actual calendar dates. This is a retrospective conditional association, with factor construction fitted over the observed period.

Table 1: Employment entry and unemployment entry twelve months later

Specification or population	Estimate (pp)	95% interval	Partial R^2 (%)	Months
A. Broad nonparticipants, common months				
Current rate and one lag	0.080	[0.032, 0.124]	10.14	343
Add calendar-month effects	0.094	[0.038, 0.148]	12.11	343
Current rate and twelve lags	0.090	[0.038, 0.138]	11.38	343
Month effects and twelve lags	0.096	[0.025, 0.157]	12.62	343
B. Restricted populations, month effects and twelve lags				
Explicit no to wanting work	0.147	[0.013, 0.276]	4.72	343
Ages 25–54	0.034	[-0.030, 0.091]	0.51	343

Notes: All rows include the common employment-entry factor and current unemployment entry. Panel A uses identical months and the same indicator standard deviation, fixed at the short-history twelve-month value. Panel B refits the indicator for each population and uses its own point-sample standard deviation. Twelve lags means the preceding twelve monthly unemployment-entry rates in addition to the current rate; the first two rows include only one preceding rate. Calendar-month effects enter the regression. Intervals are pointwise percentiles from 2,000 construction-refitting bootstrap draws with 24-month blocks. With 18-month blocks, the explicit-no interval includes zero; the four broad-population intervals remain positive. Partial R-squared is the percentage of variation left after the other regressors that is accounted for by relative employment entry.

Our comparisons hold the sample and scale fixed so that changes in the controls have a clear interpretation. The table uses 343 common months and a fixed broad-group residual standard deviation from the twelve-month current-and-one-lag sample. Population checks use their own standard deviations. Partial R^2 measures the fraction of variation left unexplained by the benchmark that adding η_t accounts for, with all other controls and dates unchanged.

Uncertainty calculations allow the estimated common movement to vary and preserve dependence across nearby months. Each of 2,000 moving-block bootstrap draws refits standardization, principal components, residuals, and regressions using twenty-four-month blocks; eighteen-month blocks provide a sensitivity check. Each origin month retains its original-calendar target and controls, and monthly survey rates remain fixed inputs. We report percentile intervals. The horizon figure also shows a separate simultaneous 95% band within each specification over months one through eighteen, using the maximum standardized absolute bootstrap deviation. Each horizon uses its eligible calendar dates.

4 Results

Stronger relative employment entry is associated with more unemployment entry a year later, even after the seasonal and dynamic controls. Table 1 shows estimates between 0.080 and 0.096 percentage points across the four control sets on identical dates. With both additions, a one-standard-deviation increase corresponds to 0.096 percentage points, with a pointwise interval of [0.025, 0.157]. This is about 7.0% of the mean monthly unemployment-entry

probability of 1.369%. The comparison concerns the monthly rate twelve months later, not the probability of entering unemployment at any time during that year.

The employment margin captures part of subsequent entry that recent unemployment entry leaves unexplained. Adding it raises ordinary R^2 from 66.93% to 71.10%. The 4.17-percentage-point gain accounts for 12.62% of the benchmark's residual variation, adding information to an already informative within-sample description. Economically, two months with similar unemployment-entry histories can differ in their later entry rates in a way associated with current job entry. The positive relationship links successful entry into work at one date with entry without a job later. The two destinations thus convey related, but distinct, information about the participation margin.

Seasonal controls change the economic interpretation of the horizon profile. Figure 1 shows negative estimates at months one through three in the parsimonious specification, but positive point estimates at all eighteen horizons with the combined controls. Month effects alone also make the first three estimates positive on unchanged dates. The short-horizon reversal is sensitive to seasonal controls, while the twelve-month association persists. The combined twelve-month simultaneous interval is [0.004, 0.188] and remains positive with eighteen-month blocks. These are associations between monthly populations: the figure describes the timing of aggregate entry margins.

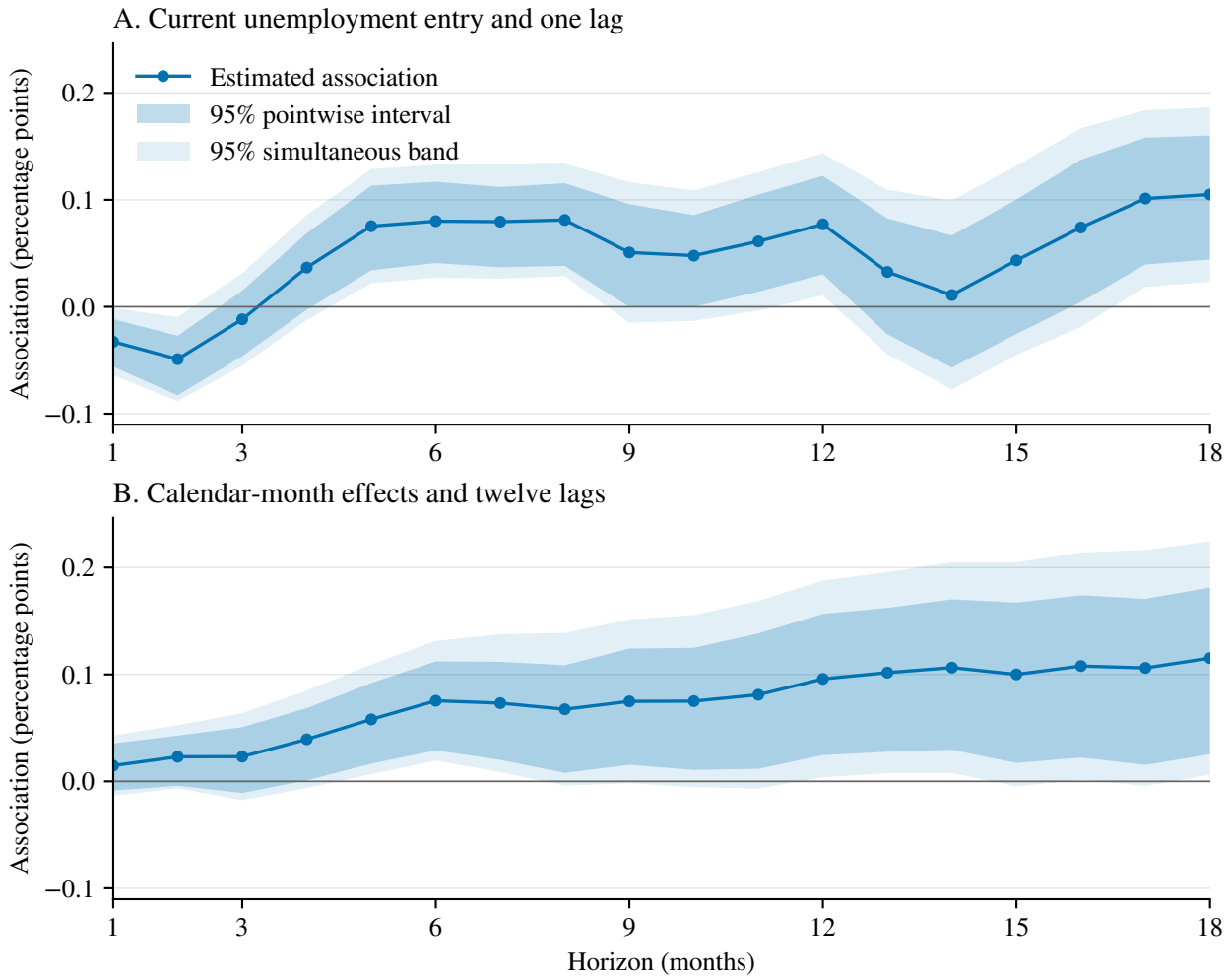


Figure 1: The horizon profile with seasonal and richer dynamic controls

Notes: Percentage points per fixed broad-population reference standard deviation. Panel A includes the common factor and current unemployment entry plus one lag. Panel B adds calendar-month effects and extends the history to twelve lags. Dark shading shows pointwise 95% percentile intervals; light shading shows a separate simultaneous 95% band over eighteen horizons in each panel. Both use 2,000 draws with twenty-four-month blocks. Native samples contain 357–373 months in Panel A and 337–349 in Panel B.

5 Conclusion

Employment entry provides a distinct perspective on how nonparticipants connect with the labor force. For the broad population without a recorded yes to wanting a job, stronger relative job entry is associated with more unemployment entry twelve months later, beyond seasonality and a year of unemployment-entry history. The economic implication is that movements into jobs and into unemployment contain complementary information about participation dynamics. Future research could examine how demographic composition and local labor-market conditions shape the connection between the two margins.

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